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Working Paper 00-05
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Grandmothers and Granddaughters: The Effects of Old Age Pension on Child Health in South Africa

Esther Duflo

Massachusetts Institute of Technology*

March 2000

Abstract

The idea that redistribution can be efficiency-enhancing if it has positive effects on human capital accumulation is often invoked to support redistribution policies in developing countries. Yet there is little evidence on whether income transfers improve human capital outcomes. This paper studies the impact of a cash transfer program on child health in South Africa. In the early 1990s, the benefits and coverage of the South African social pension program were expanded for the black population. About one-third of black South African children under age 5 live with an elderly person. This reform provides a unique opportunity to estimate the impact of an increase in cash transfers on child health, as well as the differences in impact due to the gender of the transfer recipient. Estimates suggest that pensions received by the maternal grandmother had a large impact on the anthropometric status of children, girls in particular. In contrast, I found no similar effect when a man is the pension recipient, or when it is received by the paternal grandmother.

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1 Introduction

Whether redistributive policies affect efficiency is central to the debate about the efficacy of redistribution. In particular, the case for redistribution is substantially stronger in those instances where it is shown to be efficiency-enhancing.

Redistribution can potentially be a source of efficiency gains if income transfers to the poor lead to improvements in human capital. Child nutrition is a potential conduit for such income effects, either because households are credit constrained, or because child health is in part a normal good consumed by the household. There is evidence that inadequate nutrition during childhood (and even *in utero*) affects long term physical development (Barker (1990)), as well as the development of cognitive skills.[1] This in turns affects productivity later in life (see Dasgupta (1993), Strauss and Thomas (1998), and Schultz (1999) for reviews of the evidence of the relationship between health and productivity). Therefore, low levels of investment in child health have far-reaching consequences on economic growth, distribution, and welfare. If it were shown that cash transfer programs to the poor affected child health, it could form the basis of a efficiency argument in favor of redistribution.

A related question is whether the redistribution should be in the form of cash or in-kind goods and services (cheaper access to health services or food prices subsidies, for example). Cash transfers are easier to administer.[2] Yet they are rarely used in developing countries. One reason for this is the suspicion that cash transfers could be spent by the beneficiaries in ways that would not increase efficiency. For example, parents may not spend the additional income on additional food for their children, but might instead reduce their labor supply or increase their own consumption. This suspicion is supported by the fact that the cost of meeting a child's basic nutritional requirements is extremely low, even for a poor household in a developing country.[3]

[1] Balazs, Jordan, Lewis and Patel (1986) review the biomedical and empirical literature on the relationship between early childhood nutrition and the development of intelligence. Miguel and Kremer (1999) show that school attendance is higher among children treated against worms.

[2] Of course, there is the possibility of corruption, but corruption in the delivery of services and the administration of subsidies is widespread as well. In fact, there are reasons to think that a cash transfer program is easier to monitor than a health service. To begin with, there is no way the intended beneficiaries can better verify whether they receive what they are entitled to.

[3] This is the basis of the argument made by Srinivasan (1994) against the nutrition-based poverty trap model of Dasgupta and Ray (1986).

Evidence on the effect of cash transfers on human capital accumulation in developing countries is essentially non-existent. In the United States, the evidence suggests that monetary transfers to the poor have very little impact on child human capital (Currie (1995), Mayer (1997)), and that in-kind transfers are more beneficial to children. However, the effects of parental income and monetary transfers on child outcomes are likely to be more significant among poor households in developing countries. The ideal experiment to answer these questions would be to allocate a grant randomly to some households and not to others and to compare child health in both types of households. In the absence of evidence from such an experiment, this paper exploits policy induced variations in combination with statistical modeling.

This paper evaluates the effect on child health of one of the few successful transfer programs to poor households in the developing world, the South African Old Age Pension program, a universal, non-contributory, age-tested and means-tested scheme. Historically, the Old Age Pension was racially discriminatory. At the end of the Apartheid era, the government committed to achieving parity of benefits and eligibility requirements for Whites and Blacks. This was achieved mostly by raising the benefits received by the Africans.[4] The current system is universal and non-contributory. All women above 60 and men above 65 are entitled to benefits, subject to a means test. In 1993, 80% of African women above 60 and 77% of African men above 65 received the pension. In most cases they received the maximum amount of 370 rands per month, which is roughly equivalent to twice the median income per capita in rural areas. Case and Deaton (1998) have shown that the program was effective in transferring money predominantly to poor households and, in particular, to households where poor children live.

The data used in this paper is a single cross section of 9,000 South African households undertaken in 1993. Pension recipients often live in extended households, with their children and grandchildren.[5] Close to one-third of African children under the age of 5 live with a pension recipient. Children who live with a pension recipient have, on average, relatively disadvantaged backgrounds. At the end of their study of the targeting of the pension program and its consequences on household expenditures, Case and Deaton (1998)compare height for age of children

[4]In what follows, I generally use the official denominations of racial groups in South Africa (Africans, White, Coloured and Indians). I sometimes use "Blacks" instead of "Africans" to refer to black South Africans.

[5]This is due, in particular, to Apartheid rules which prohibited the families of migrant workers, those working in the mines or as domestic servants, to join them.

in households who receive a pension and household who do not (using the same data set), and find that the difference is significantly *negative*: children who live with a pension recipient tend to be smaller for their age than other children. As they note, this is not surprising, because the rapid expansion only began in 1991, and the pension was fully operating in all areas for less than a year. Child height is a stock, which reflects accumulated decisions since the birth of the child. Children who grew up in otherwise less favorable environment are, therefore, expected to be smaller even if the pension affected current nutrition flows. Case and Deaton acknowledge that the effect of the pension program cannot be estimated from a simple comparison of households where there is a pension recipient and other households. They do not suggest an alternative estimation strategy.

To estimate the effect of receiving a pension on child nutrition, this paper exploits the fact that height reflects accumulated investments in child nutrition. The larger the proportion of her life during which a child was well nourished, the taller she will be, given her age. Due to the expansion of the program in the early 1990s, those of qualified age became more likely to receive a pension. The benefits also became substantially larger. Thus, children born after the expansion of the program are more likely to have spent a larger fraction of their lives well-nourished, if they live with a pension recipient.

The basic idea underlying the identification strategy is, therefore, to compare the differences in height between young children and older children across households, according to whether or not there is an age-qualified person living in the household. This strategy allows height for age to vary systematically with age and with eligibility status of the household. The identifying assumption is that any difference between children in eligible and non-eligible households would have been the same in all age groups in the absence of the program. The identification assumption may not be satisfied if households where an eligible person lives are also targeted by other programs during the same time period. I investigate this issue by examining whether the results are sensitive to alternative definitions of the control group (to include, for example, only non-eligible three generation households). Further, an implication of the identifying assumption can be directly tested. Weight for height is a short run measure of well being, which should be affected by the increase in income for *all* children (not only for the younger ones). Therefore the difference between the weight for height of eligible children and that of non-eligible children should be the same among young and old children.

It is straightforward to extend this strategy to estimate separately the effects of pensions received by women and by men. The South African experience is therefore a unique opportunity to determine whether the gender of the transfer recipient matters. Even if households with female recipients differ systematically from households with male recipients, this will be taken into account by the identification strategy, as long as these differences are additive with respect to the age of the child. Likewise, the effects can be separately estimated according to the gender of the 'middle generation', i.e. the parent of the child (who is the son or the daughter of the eligible member).

Taken together, the results in this paper can, therefore, contribute to answering the two questions we asked. Can cash transfers lead to improvements in child health ? Are they likely to be consumed by other household members? The evidence suggests that the pension had a beneficial effect on girls' nutrition (less so on that of boys) when it was received by a woman, and when that woman was the maternal grandmother, but not in other cases. This suggests that cash transfers can, in some circumstances, have a positive effect on child health. However, it also shows that some transfers do not benefit children at all, even when the transfers are large and reach households where children live.

The remainder of this paper is organized as follows. The next section presents the theoretical and empirical background on the relationship between household income, social transfers and child nutrition. Section 3 presents a brief history of the South African Old Age Pension program. In section 4, I present the identification strategy and the main results. In section 5, I discuss potential problems of the identification strategy, and I present additional evidence to address them. In section 6, I present and interpret further results: the effects of the pension in various subsamples, a comparison of the impact of pensions received by men and of those received by women, and a comparison across gender in the middle generation (i.e. the gender of the parent of the child whose grandparent receives the pension). Section 7 concludes.

2 Background: the relationship between Income and nutrition

2.1 Conceptual background: the relationship between income and nutrition

Two main models explain why an increase in income can affect child height. They emphasize different aspects of the question, and do not exclude each other.

First, in a pure investment model, healthy individuals are more productive, and this productivity will be rewarded in self employment (Strauss (1986)), in the labor market (see Rosenzweig and Schultz (1982), Schultz (1996)) or in the marriage market (Foster and Rosenzweig (1999)). Parents invest in child health in anticipation of these returns. In this model, a change in income will change the level of investment into child human capital only if households are constrained in their ability to invest in child health by their inability to borrow. Any relaxation of the credit constraint (even if this does not correspond to an improvement in permanent income) will then result in better child nutrition.

Second, child health (and, in general, child quality) can be considered as a consumption good. Household members derive utility from healthy children. If child quality is a normal good, any change in permanent income will affect child health even if households are not credit constrained. A transitory change in income, by contrast, should not affect investments into child health, unless the household is credit constrained.

Third, both models can be extended to take into account the fact that the household is a multi-person decision unit. If household members have different preferences (for example, different effective discount rates, or different valuation of child health), the way in which the preferences are taken into account in the household decision making process will matter for final outcomes (See Chiappori, (1988,1992) Bourguignon, Browning, Chiappori and Lechene (1994), Browning and Chiappori (1998)). The weight given to each individual's preference depends on 'distribution' factors. The fact that the person controls a source of income, for example, can give her more influence on the decision of how this income is allocated.

Pension income has several characteristics which might lead its effect on child nutrition to be smaller or larger than the effect of another income shock. First it is a regular income. Few Africans have a stable labor force attachment or a regular source of income, especially in rural areas. Therefore, in households where there is a pension recipient, pension benefits often are the most regular sources of income. Defining permanent income over a one-year horizon (as in Paxson (1992)), pension income is permanent, while non-pension income is in part permanent, in part transitory. The propensity to spend out of pension income might therefore be greater than the propensity to spend out of non-pension income. Moreover, anecdotal evidence (Lund (1993)) suggests that pension recipients can borrow against future pension income. She reports

that in some villages, only households where a pension recipient lives can receive credit. If the pension actually relaxes the household credit constraint, this implies that the propensity to spend on child health out of pension income should be especially high.

Second, and however, if we consider a larger horizon of permanent income, a rand of pension income today represents less than a rand of permanent income, since it is tied to an elderly person and will stop when the elderly person dies. If households are not credit constrained and child quality is a normal good, the propensity to spend on child health should be smaller out of pension income than out of non-pension income. A pension received by a man and a pension received by a woman differ also in this respect. On average, men receive the pension for a much shorter length of time because they not only are eligible later in life but also have shorter life expectancy. Therefore a rand of pension income received by a man represents less, in term of permanent income, than a rand of pension income received by a woman.

Third, pension income is received by an elderly person (and more often by a woman than by a man). This improves her bargaining position, and might influence the allocation of this income. If the elderly person has different preferences than the younger members of the household, the pension will have a different effect on children than any other source of income. The middle generation (the child's parents) might also play a role, if the pension recipient transfers part of her bargaining power to her child, rather than to her child's spouse. The last section of this paper is largely devoted to these issues.

Case and Deaton (1998) conclude that "a Rand is a Rand" for most expenditures, and that pension and non pension income seem to be spent in similar ways. However, the expenditure survey does not pick up distinctions in how the expenditures (within each category) are allocated between household members. It is therefore possible that child health responds differently to pension and non pension income even if expenditures do not.

2.2 The relationship between income and nutrition: empirical literature

There is a body of evidence showing that height is positively correlated with household resources in developing countries (see Thomas, Strauss and Henriques (1990a,1990b), Thomas, Lavy and Strauss (1992), Sahn (1990)). A common difficulty faced by this literature has been identifying exogenous sources of variation in income. Income and child heath are jointly determined. In a static household production model, non labor income or assets might be treated as exogenous.

However, in the more realistic dynamic model, assets and income are endogenous. Moreover, they are difficult to measure accurately, and, even if we accept the static model, they are in general not randomly assigned. Households which have assets and non-labor income tend to live in less crowded areas with better hygiene, cleaner water and better health care services. For example, Thomas, Strauss and Henriques (1990a) find that controlling for observed community characteristics reduces the estimated effect of household per capita income by half, and makes it insignificant. Most estimates of the effect of family resources on child health are therefore likely to be biased by the omission of unobserved family or community characteristics. Therefore, this evidence does not provide much guidance on whether cash transfers to poor households could improve child health. Foster (1995) focuses explicitly on credit market imperfections and finds that the weight of children living in households with restricted access to credit fluctuated more as a result of the 1988 flood in Bangladesh than that of other children.

A smaller literature tests whether income in the hands of the women of a household has a different impact on intra-household allocations than income in the hands of the men. The evidence suggests that, compared to income or assets in the hand of men, income or assets in the hands of women are associated with larger improvements in child health (Thomas (1990)), and larger expenditure shares on household nutrients, health and housing (Thomas (1993)). However men's and women's income are not exogenous, and non-labor income (used in Thomas (1990,1993)), will not be a valid instrument in general. In addition, if part of a woman's assets were given by her family before or at the time of her marriage, the marriage market will in general insure that her assets and non-labor income will be correlated with (potentially unobserved) characteristics of her husband. This will invalidate any comparison, since the coefficient on the wife's non labor income will pick up the effects of the husband's unobserved characteristics. Behrman (1997) concludes his review of the empirical literature on the remark that the available evidence does not allow to reject income pooling. Lundberg, Pollak and Wales (1996) address these difficulties by studying a reform of the child maintenance in the United Kingdom. Prior to 1977, benefits were in the form of a tax allowance (and amounted to an increase in the take home pay of the man). After 1979, a child benefit was paid out directly to the child's mother. They show that, in families with children, the ratio of women and children's clothing to men's clothing became larger after the reform.

In the United States, some studies have tried to identify plausibly exogenous sources of

variation in parental income (in particular, monetary transfers) and to measure their impact on child outcomes. Shea (1997) studies whether children's outcomes (education and subsequent labor earnings) are correlated with their father's union status, job loss, or industry and finds no effect of these variables on child outcomes, except among the poorest families. Mayer (1997) and Currie (1995) examine the impact of several transfer programs on children welfare. Both find little or no effect of monetary transfers such as AFDC, but Currie (1995) finds that in-kind programs (such as Medicaid, feeding programs, and Head Start) have positive impacts on children welfare. The effects of parental monetary transfers on child outcomes are likely to be of a larger magnitude among poor households in developing countries. However, in developing countries, most social programs are in-kind, and there are almost no instances of monetary transfers to the poor, and, therefore, no evaluation of their consequences. The South African Pension program provides an unusual opportunity to evaluate the possible effect of such a monetary transfer.

3 Description of the program

3.1 The South African Old Age Pension program

This section present a brief history and an overview of the functioning of the South African Old Age Pension program. It draws extensively from Van der Berg (1994), Lund (1993), and Case and Deaton (1998).

Social pensions were first introduced in the 1920's for Whites. They were intended mainly as a social safety net for a minority of white workers who were not covered by occupational pensions. The pensions were gradually extended, but with very dissimilar benefits levels, to other race groups. During the Apartheid era, the system was racially discriminatory in many respects. First, different means tests were applied to each race group. In 1984 for example, benefits were withdrawn for incomes larger than R 700 per annum (R 500 in Kwa Zulu) for Blacks, and for incomes larger than R 2250 per annum for Whites. Second, the benefits levels were different. In the early 1980s, benefits for Whites were 10 times higher than those for Blacks.[6] Third, the delivery systems were different. Whites' pensions were distributed through

[6] Note that non-pension incomes of Africans were also much smaller, therefore as a fraction of income, the difference was much smaller (Van der Berg (1994)).

the postal offices, while Africans' were distributed through mobile pay points that did not go very far out into rural areas. Finally, officials often intentionally underestimated the age of the person, took people off of the computer lists, or otherwise limited the access of legally eligible Africans in order to save on the cost of pensions (Lund (1993)).

Pressures for equity in the treatment of racial groups became strong towards the end of the Apartheid. In 1989, the government committed to achieving racial parity in pension treatment (Van der Berg (1994)). The extension of the social pensions to the whole population took several years and was operating fully in all areas only by the beginning of 1993. The benefits for Africans improved gradually in the 1980s (from 1555 1990-rands per year in 1980 to 2096 1990-rands per year in 1990), while those for Whites declined rapidly), and then much faster in the 1990s. Table 1 (taken from Van der Berg (1994)) shows the maximum annual pension payable from 1987 to 1993, expressed in 1990 rands. Benefits were roughly constant from 1987 to 1990 and started increasing rapidly in 1991. Benefits were multiplied by 1.5 between 1990 and 1993. The maximum benefit in 1993 (the survey year) was 370 rands ($3 per day). In comparison, the mean monthly household income per capita of Africans in the sample was 149 rand in 1993 (the mean monthly income per adult was 404 rands). Given the high level of unemployment, the pension recipient is frequently the main income earner in the household. As a fraction of income, the pension program had become a unusually generous transfer program by 1993.

In 1992, the means test was modified, and unified across race. The current system is universal and non-contributory. Payments are made to women older than 60 and to men older than 65, subject to a means test. For couples, household resources are roughly divided by two when calculating the means test. Importantly, the income of other members of the household is not taken into account when implementing the means test. There are, therefore, no direct incentives to partition the household or to stop working for other household members. In practice, the means test does not seem to be applied very finely. It is mainly effective in excluding most Whites and those Africans who already have a private pension.

In 1993, 80% of the Africans eligible on the basis of their age were receiving a pension. Of those, most receive the maximum amount. There is no good estimate of the coverage earlier on. First, social pensions were administered by several different administrations, which made any evaluation difficult. Second, surveys (including the 1991 census) excluded the "independent homelands", where many Africans live. The coverage increased substantially in the 1990s, due to

a new attitude within the administration, a modification of the means test, the computerization of the system, and substantial improvements in the delivery system.[7]

3.2 Data and descriptive statistics

The data for this paper came from the national survey of South Africa carried out jointly by the World Bank and the South African Labor and Development Research Unit at the University of Cape Town. During the last five months of 1993, 9,000 randomly selected households from all races and areas, including the "independent homelands", were interviewed. This is a multipurpose household survey similar to most World Bank Leaving Standards Measurement surveys. As part of the survey, measurements of the height and weight of all children aged less than seven years were taken. Environmental factors are especially important determinants of child height in early childhood. Therefore, the World Health Organization recommends limiting the analysis of height measures to children 0 to 5 year old (WHO (1986)). In addition, there appears to have been difficulties in the measurement of the oldest children.[8] I, therefore, follow previous studies (Case and Deaton (1998),Le Roux (1995)), and restrict the sample to children aged 6 to 60 months.

Descriptive statistics of the sample of black children are presented in table 1. Columns 1 and 2 show the means of the variables used in the analysis by pension status. Households where there is a pension recipient tend to be poorer than the mean (their pre-pension monthly income is 661 Rands, compared to 930 Rands in other households). Income after pension is higher in households that receive a pension, but income per capita of pension recipient households remains slightly smaller. They are, not surprisingly, often characterized by the presence of a grand parent (93%), and the absence of the child's father (64%), or mother (17%). They are also more likely to live in a rural area.

Even if child nutrition has improved as a result of the extension of the coverage and benefits of the Old Age Pension program, height given age still reflects past deprivations or illnesses, especially among the oldest children. Height is strongly associated with age, so even small vari-

[7] For example, in the province of Kwa Zulu Natal, the pension is distributed once or twice a month through mobile pay points equipped with ATMs with fingerprint recognition system (Case and Deaton (1998)).

[8] Some six and seven year old children were recorded by the interviewers to be eight, and thus were not measured. It seems very likely that if the child was tall the surveyor would have assumed the child was older, and therefore the surveyor would have mistakenly excluded that child.

ation in age composition will make any comparison meaningless. I follow the norm recommended by the World Health Organization and applied by most researchers: for each age in months, I construct height for age Z-scores by subtracting the median and dividing by the standard error in this age and sex group in the NCHS reference population (a group of well-nourished American children).[9] Descriptive statistics of height for age and weight for height are presented in panel C. Children are smaller, controlling for their age, in household where there is a pension recipient. Interestingly, weight given height (a measure of short run nutritional status) is larger in eligible households.

Family background, which varies systematically with pension receipt, directly affects nutrition. Therefore, the comparison of anthropometric indicators or expenditures on child health and nutrition across pension and non-pension households would not be an unbiased indicator of the effect of pension income on child health, as noted by Case and Deaton (1998). In the next section, I propose an empirical strategy to overcome this difficulty of identification.

4 Estimating the effect of Social pension on Child health

4.1 Identification strategy

In practice, in developing countries, human growth deficits are caused by two preventable factors, inadequate food and infections. Genetic factors matter for child height, but they become more critical in adolescence. In childhood, height for age and weight for height Z-scores are widely considered to be "the most useful tool for assessing the nutritional status of children" (WHO (1986)).

Height given age of young children depends on accumulated investments over the life of the child (Martorell and Habicht (1986)). I capture this in the following formulation:

$$h_i(a) = \overline{h(a)} + \epsilon_{ai} + \phi(\frac{N_{0i}}{N^*_{0i}}, .., \frac{N_{ai}}{N^*_{ai}})\overline{\sigma_a}, \quad (1)$$

where $h_i(a)$ is the height attained by child i at age a, $\frac{N_{si}}{N^*_{si}}$ (for $s = 0$ to a) is the ratio of the nutrition and other necessary inputs (primary health care,...) received by the child relative to what would be optimal, $\overline{h(a)}$ is the average height of well nourished children of this age, $\overline{\sigma_a}$ is the

[9] Note that this normalization does not affect the analysis, which relies on the comparison of height between from eligible household and those of the same age from non-eligible households

standard deviation of well nourished children of this age. The function $\phi(.)$ is weakly increasing in all its arguments and $\phi(1,1,..,1)$=0.[10] This formulation captures the idea that the child has an ideal height at a given age, which is genetically determined $(\overline{h(a)} + \epsilon_{ai})$, and that inadequate nutrition over her whole life prevents her from achieving it. It reflects the standard practice of expressing height for age in the number of standard deviations from the average height in the reference population, by expressing the distance from the ideal height in numbers of standard deviations in the reference population.

Some properties of the function $\phi(.)$ are documented in the medical literature. First, nutrition at a very early age (in the womb and in infancy) has long lasting consequence on child height, and in fact, on adult health as well (Barker (1990), Scrimshaw (1997)). Second, the possibility of catch up skeletal growth after an episode of low growth in infancy is limited, but weight for height recovery is faster.[11] Most stunting and catch up occur between 6 months and 24 months of age. Stunting after 24 months of age generally reflect the interaction of nutrition and infection at prior ages (Scrimshaw, personal communication, Martorell and Habicht (1986)).

The amount of nutrition (and, more generally, health inputs) a child gets is in turn determined by family decisions. As in most studies of the determinants of nutritional status, consider a household production model in the tradition of Becker (1965). Household members exercise choices over consumption (including leisure), and the number and the quality of surviving children. The resulting maximization problem generates reduced form expressions of height for age and weight for height as functions of individual, family and community characteristics (unobserved or observed) and exogenous income sources of family members. The decision to apply for a pension should be considered as endogenous. I, therefore, focus on age-eligibility for the pension. Ignoring the possibility of endogenous family formation (we will come back to this issue below), I write the reduced form equation for child nutrition at each date as follows:

$$\frac{N_{ti}(d)}{N^*_{ti}} = \lambda E_i * g_i(t) + u_i + \nu_{itd}, \tag{2}$$

[10]This formulation captures the situation in a developing countries, where most arguments of the $\phi(.)$ function will be below 1.

[11]For example, a study of Jamaican children (Ashworth (1969)) shows that children's weight for height recovers quickly from episodes of acute malnutrition, but that once normal weight for height is achieved, the body stops accumulating nutrienst that would allow it to achieve faster skeletal growth.

where t is the date considered, d is the date of birth of the child, $E_i = 1$ if the family is eligible and $g_i(t)$ indicates the value of eligibility (for the household) in terms of money. It reflects both whether the household receives the pension and its amount. The evidence on the program expansion presented in section 2 suggests that, if the household was not forward looking, $g_i(t)$ was more or less flat from 1988 to 1990 and increased rapidly from 1991 on. Note that if the expansion of the program was anticipated, the pension could affect even older children, provided that the household (or eligible member who would receive the pension) does not face credit constraints.

The term u_i in this equation captures family background, which directly affects all children's nutrition. This term is correlated with eligibility status, since demographic composition both reflects and affects preferences and endowments of a family. Households which have an eligible member tend to be poorer, larger households, with at least one parent more likely to be absent. This precludes inference of the effect of the pension from comparisons of the nutritional status of children in eligible and non eligible households.

However, since all children were measured around the same date, if the nutrition is indeed affected by the pension, older children have had longer periods of low nutrition. Using equation 1, we see that if the possibility of catch up growth is limited, older children, who benefited from the program for a fraction of their lives and grew up in otherwise less favorable environments, should be smaller in eligible households than in non-eligible households. However, if the pension had an effect, the difference between eligible and non eligible children should be smaller for younger children, or may even be reversed.

The basic idea of the identification strategy is thus to compare the differences between the height of children in eligible and in non eligible households and between children exposed to the program for a fraction of their lives and children exposed all their lives. The descriptive statistics of height and height for age in different sub-samples presented in tables 3 illustrate this identification strategy. Columns 3 and 4 show the means of height for age among households where there is an eligible woman and an eligible man, respectively. Column 5 show these means among households where there is no eligible member. Among children born before January 1992, both boys and girls are smaller in households where there is an eligible member (woman or man) than in other households. However, girls born after January 1992 are taller if they live

with an eligible member (especially with a woman). This is not true for boys. This suggests that the pension had an effect on the nutrition of girls.

I present a non-parametric version of this comparison in the next subsection.

4.2 Non-parametric approach

4.2.1 Non-parametric regressions

The least restrictive implementation of the identification strategy is to plot height or height for age as a function of date of birth in eligible and non eligible households, and to examine the relative positions of these two curves. This non-parametric approach is compelling for two reasons. First, it took some years to achieve universal coverage and parity in the benefits. Second, we have little knowledge about the functional form of the child health production function (the function $\phi(.)$). However, we know that the coverage and the benefits were increasing over the period (so that young children were more exposed to the program than older children) and that there is no complete catch up growth.

In practice, I estimate non parametrically the equations $h_E^J(d) = \psi^E(d) + \epsilon_i$ (where d is the date of birth of the child) in the sample of eligible households and $h_i^N(d) = \psi^N(d) + \epsilon_i$ in the sample of non eligible households. We are interested in the shape of the function $dh(d) = \psi^E(d) - \psi^N(d)$.

Figure 1-A show non parametric (Fan's (1992) locally weighted) regressions of height in centimeters as a function of date of birth, in eligible (straight line) and non eligible (broken line) households. Not surprisingly, height is strongly related to age. The interesting point is that the curves have the relative positions predicted in the preceding discussion. Children born before January 1992 are smaller in eligible households. Children born after January 1992 are taller in eligible households. Because in this picture any difference is swamped by the age profile, I present in figure 1B the difference between this two curves. The difference is negative, and roughly stable, for all children born until January 1991. Then the gap starts closing, until the difference become positive around January 1992. The difference increases until July 1992, and then it stabilizes.

In figure 1-C, I show this difference for boys and girls separately. Each of the curve has roughly the pattern discussed above, but among elder children, the handicap of eligible boys

over non eligible boys is smaller than the comparable difference for girls, and among younger children, their advantage is also smaller than the comparable difference for girls. This suggests a larger impact of the program on the nutrition of girls. It might also suggest, more generally, that the effect of background and income on the nutrition of boys is smaller than on that of girls. Eligible boys suffer less from growing up in generally poorer families, and non-eligible boys suffer less from not receiving the transfer. This result echoes, for example, the finding in Rose (1999) that, in rural India, girls suffer more from adverse shocks than boys do.

Figure 2 shows non parametric regressions of height for age Z-scores as a function of date of birth, in eligible (straight line) and non eligible (broken line) households. These curves have the shapes traditionally found in developing countries: height for age Z-scores decline fast in the first two years of life, and then stabilize. The relative position of the two curves is the same as in figure 1-A (but the pattern is much clearer because height for age does not vary as much with age as height does). This pattern is even more striking when looking at girls alone (figure 2-C). The advantage of girls living in eligible households over those who are not, among young girls, is as large as their handicap among older girls.

This evidence suggests that the increase in pension contributed to improve child nutrition, especially that of girls, and resulted in faster growth for the youngest children. Not only did the pensions help the girls living in an eligible household to bridge the gap with girls living in non-eligible households, but also they seemed to have help them to do *better.*

4.2.2 Average derivatives

An easy way to summarize these patterns and to evaluate whether they are significant is to calculate estimates of the average derivative of the function $dh(d)$ over the relevant ranges. To do so, it is convenient to estimate the functions ψ^E and ψ^N with a cubic splines formulation. In practice, I simply estimate OLS regressions of height (or height for age) on a cubic polynomial in the date of birth (expressed in months, with January 1990 defined as month 0) and four terms defined as $\mathbf{1}_{(d\geq l)}(d-l)^3$ where $l = -12$, 0, 12 and 24 respectively and $\mathbf{1}_{(d\geq l)}$ is a variable indicating whether d is greater or equal than l. I then calculate the analytical expressions for the estimated function $\widehat{dh(.)}$ and its derivative. I can then compute derivative over any chosen range, as well as its standard error.

In table 4 (columns 1 and 2) I present estimates of the average derivative of the function

$\widehat{dh(.)} = \widehat{g^E}(.) - \widehat{g^N}(.)$ for children born between 06/88 and 12/90 (column 1) and for children born between 01/91 and the 06/93 (column 2).[12] The break point in December 1990 is justified by the pattern of benefits, which increased in levels starting in 1991. In panel A, I show the results for height in centimeters, for boys and girls separately. For boys, the average derivative is not significantly different from 0 in both subsamples. For girls, the average derivative is close to 0 in the first period, and significantly different from 0 in the second period. Similar results are obtained for height for age Z-scores. Eligible girls gain on average 0.093 centimeters, or 0.026 standard deviations each month relative to non eligible girls. The point estimates of the average derivative in the young sample is twice as large for girls as it is for boys, although the equality of the two coefficients is not rejected.

4.3 Estimating the impact of pension on nutrition

This suggests that the program had an impact on the nutrition of girls. In order to perform comparisons with other programs, we would like to be able to give an estimate of the effect of a rand of pension on child nutrition.

This requires to place some additional restrictions on the shape of the function linking nutrition and child height defined in equation 1. The following weighted average formulation for the function $\phi(.)$ corresponds to the discussion of the medical evidence:

$$\phi(\frac{N_0}{N_0^*}, .., \frac{N_a}{N_a^*}) = \alpha \frac{N_a}{N_a^*} + (1-\alpha)\frac{\sum_{s=0}^{a} \delta^s \frac{N_s}{N_s^*}}{\sum_{s=0}^{a} \delta^s} - 1, \quad (3)$$

where α and δ are parameters smaller than 1. The parameter δ captures the relative importance of early childhood nutrition, while α captures the possibility of catch up growth during phases of good nutrition.

Denoting $g(t)$ the average of $g_i(t)$, taken over i the average monetary value of eligibility in the sample), we can use equation 2 to express the average of $\frac{N_{si}}{N_{si}^*}$ among eligible households: $E(\frac{N_{si}}{N_{si}^*}|E_i = 1) = \lambda g(t) + \overline{u^E}$, and non eligible households: $E(\frac{N_{si}}{N_{si}^*}|E_i = 0) = \overline{u^N}$. All the children where measured around the same date (denoted T), therefore $a = T - d$. Using equation 1 we can can now express the difference between the average height of children in eligible and non eligible family, as a function of age.

[12] I also restrict the sample to children aged 4 to 60 months

$$\frac{\overline{h^E(a)} - \overline{h^N(a)}}{\overline{\sigma_a}} = \alpha\lambda g(T) + (1-\alpha)\frac{\sum_{t=d}^{T}\delta^{t-d}\lambda g(t)}{\sum_{t=d}^{T}\delta^{t-d}} + \overline{u^E} - \overline{u^N}. \tag{4}$$

The function $g(t)$ is increasing. Therefore, if α is not equal to 1 and δ is not close to 0, the difference should be increasing over all ages. However, the figures and the average derivatives presented in table 4 suggest that the difference was stable for all children born up to 01/91 and started increasing only subsequently. This suggests two things: first, δ is in fact close to 0, and second, the function $g(t)$ is closely linked with the actual expansion of the program. Because the program expansion was anticipated, this suggests that future pension recipients cannot borrow against their benefits.

If δ is equal to 0, the formula reduces to:

$$\frac{\overline{h^E(a)} - \overline{h^N(a)}}{\overline{\sigma_a}} = \alpha\lambda g(T) + (1-\alpha)\lambda g(d) + \overline{u^E} - \overline{u^N}, \tag{5}$$

Assuming that δ is equal to 0 (which seems to be consistent with the non-parametric evidence), we can implement a simple difference in differences estimator, using height for age Z-scores as the outcome variable. Consider a group of children born after the full expansion, and a group of children born before the beginning of the expansion. The difference of equation 5 evaluated for the younger children (Y) and equation 5 evaluated for the oldest children is equal to $(1-\alpha)\lambda(g(Y) - g(O))$. $g(Y)$ is known (up to the choice of a functional form for $g(.)$; an average value for $g(O)$ must be obtained from data about the program before the expansion; finally assumptions must be made about α to recover λ.

4.3.1 Reduced form: effects of eligibility

The discussion in the preceding subsection suggests the following formulation:

$$h_{ifk} = \pi \mathbf{1}_{(k=1)} * T_f + \beta T_f + \sum_{j=1}^{3}\gamma_j \mathbf{1}_{(k=j)} + X_{ifk}\delta + \sum_{j=1}^{3}\mathbf{1}_{(k=j)} * X_{ifk}\lambda_j + \epsilon_{ifk}, \tag{6}$$

where h_{ifk} is the height for age Z-score of a child born in cohort k, in family f. T_f is an indicator variable equal to 1 in families where there is at least one eligible member, and to 0 otherwise. In 20% of eligible households, there are two eligible members (in only 3 households, there are three eligible members). The two eligible members are of the opposite sex in the majority of cases. Below, I will examine whether having two eligible members has a greater positive effect

than having only one, and I will also examine the impact of eligibility of both genders. $\mathbf{1}_{(k=j)}$ denotes an indicator variable equal to 1 if k is equal to j, and 0 otherwise. Children born in January 1992 or later ($k = 1$) form the most exposed group. The control group is formed of three groups: children born January 1990 - December 1990 ($k = 2$), January 1989 - December 1989 ($k = 3$), and before 1989 ($k = 4$). These children born in December 1990 or before were all born before the program started its expansion.[13]

The last two terms, X_{ifk} and $\sum_{j=1}^{3} 1_{(k=j)} * X_{ifk}$, are family background variables and family background variables interacted with cohort dummies, respectively. The identification assumption is that in the absence of the program the cohort trend would have been the same for eligible and non eligible children. This is more likely to be true if eligible and non eligible children are similar, conditional on the control variables, even among older children.[14] This makes control variables important. I therefore estimate three forms of equation 6: without control variables, with two sets of control variables non interacted, and interacted with age dummies. The first set of variables includes family background variables: mother's and father's education, rural, urban or metropolitan residence, mother's and father's age, but not family composition variables, father absent or mother absent.[15]

The second set of variables includes family size and the number of household members in the following age categories: 0-5, 6-14, 15-24, 25-49, and 50 and above. When I control for a household member aged 50 or above, I take advantage of the fact that the pension becomes available only at age 60 (for a woman), or 65 (for a man). Households with both a member above 50 and a child are similar to eligible households where there there is a child but no members over 50.[16] There may be systematic differences between them (ideally we would like to use a sharp discontinuity at the age of eligibility, but this is difficult due to mistargeting of the pension by age), but it is more likely that the trends between cohorts are similar in the two groups. In this

[13] I have also run the regression in a sample that include the children born during the increase (January 1991-December 1991) are included in the control group, and I in a sample that exclude children born in 1992, and I find very similar results in both cases.

[14] The more different they are, the more we can suspect that different things could happen to them over the period.

[15] I have replaced these variables by sample means in cases where the father or the mother of the child were absent, to avoid selecting the sample using this criterion.

[16] This is the basis of the identification in Case and Deaton (1998) and Bertrand, Miller and Mullainathan (1999) who are restricted to use pure cross-sectional variation in their outcomes of interest.

case, the identifying variation is just the age of the grandparent (interacted with the child's year of birth). Any age-varying effect of the presence of the grandparent is thus controlled for.

Equation 6 is estimated by OLS, and standard errors are adjusted to take into account the correlation of errors terms between children in the same families, as well as heteroscedasticity.

4.3.2 Effects of the pension

In 1993, 82% of the children living with an eligible woman and 79% of those living with an eligible man lived with a pension recipient. 6.5% of the households where nobody was eligible received a pension.[17] As discussed earlier eligibility status can be used as an instrument for receipt of a pension after the extension of the program.

Assume first that $\alpha = 0$ and there was no pension to speak of before 1991 ($g(O)$=0). Then we could estimate λ by estimating the following equation using two stage least squares (2SLS).

$$h_{ifk} = \lambda \mathbf{1}_{(k=1)} * P_f + \beta T_f + \sum_{j=1}^{3} \gamma_j \mathbf{1}_{(k=j)} + X_{ifk}\delta + \sum_{j=1}^{3} \mathbf{1}_{(k=j)} * X_{ifk}\lambda_j + \epsilon_{ifk}, \tag{7}$$

where P_f is a dummy equal to 1 if the household receives a pension (in 1993). The excluded instrument is the interaction $\mathbf{1}_{(k=1)} * T_f$. Alternatively, we can replace P_f by the amount of pension received by the household.

The corresponding first stage is therefore:

$$\mathbf{1}_{(k=1)} * P_f = a\mathbf{1}_{(k=1)} * T_f + bT_f + \sum_{j=1}^{3} c_j \mathbf{1}_{(k=j)} + X_{ifk}d + \sum_{j=1}^{3} \mathbf{1}_{(k=j)} * X_{ifk}l_j + v_{ifk}, \tag{8}$$

The straight 2SLS estimate is a lower bound of the effect of the average pension received on nutrition (since it assumes that $g(O)$=0 and α=0). A range of estimates can be obtained by dividing this estimate by guesses of reasonable values for α and $g(Y) - g(O))/g(Y)$.

4.4 Differences in differences formulation: Results

- **Reduced form estimates**

[17] This comes in part from the fact that I consider as eligible people who were eligible in 1992 (i.e women above 62 and men above 67) and in part from the fact that individuals receive a pension despite not being eligible. This is especially frequent among men, because in some regions, officials applied the 60 years eligibility threshold to men as well as to women.

Estimates of equation 6 are presented in columns 1 to 3 in table 5. Column 2 adds family background variables and family control variables interacted with year of birth dummies, and column 3 adds the number of family members in each age category, interacted with year of birth dummies. Results for boys are presented in panel A. Results for girls are presented in panel B. These results are consistent with the patterns displayed by the graphs.

For both boys and girls, the coefficient of the indicator for eligibility (the difference between the height for age of eligible children born before January 1992 in eligible and non eligible households) is negative, but not significant. It is similar for boys and girls. For girls, it drops from -0.16 to -0.09 when I control for the presence of a household member over 50, which indicates that the two groups are indeed made more similar by introducing this control. For girls, the lowest point estimate of the coefficient of the interaction between eligibility and being born after January 1992 is 0.44 (with a standard error of 0.27) which is more than twice as large as the coefficient of the indicator for eligibility. Young girls are taller if they live with an eligible individual than otherwise, while the reverse is true for old girls. The coefficient raises to 0.56 when I control for household members' age, interacted with year of birth dummies, and thus use only the variation in the age of members aged 50 or above as a source of variation. For boys, the effect is close to zero and insignificant in all specifications.[18]

- **2SLS estimates of Effect of the pension**

The first stage estimate (equation 8) is presented in column 1, table 6. There is a strong association between pension receipt and eligibility for pension and the first stages are highly significant in all sub-samples. The coefficient of the interaction between eligibility and belonging to the youngest cohort is 0.75, with a t. statistic above 20.

2SLS estimates of equation 7 are presented in table 7 (column 2). For comparison, the OLS effect of the pension receipt is presented in column 1 (this is the coefficient of a dummy equal to 1 if the household receive the pension). As in Le Roux (1995), I find a negative and insignificant effect of the pension, both for boys and girls, when it is estimated using OLS. In contrast, for girls, the IV estimates are positive and large (0.58 -with a standard error of 0.37-

[18]Controlling for the presence of a grandparent instead or in addition to controlling for the presence of a household member above 50 does not change the results relative to column (3). Controlling for family size, number and age composition of children does not affect the estimates in column (2). Controlling for month of birth instead of year of birth does not change the estimates either.

without controlling for the presence of the grand mother, 0.82 -with a standard error of 0.47- controlling for the presence of a household member above 50). They are small and insignificant for boys. Effects found for girls are large, but note that the pension represents a large income transfer (twice the median household income per capita in rural area). In columns 5 in table 7, I present estimates of equation 7, but where I replace an indicator variable for receiving the pension by the pension income received. Again, I find positive coefficients for girls, significant when I control for the presence of an household member above 50, and a small and insignificant coefficients for boys.

In table 8, I provide a range of estimates of λ (the effect on height for age) which are implied by these IV estimates, for plausible values of the catch up parameter α and of the probability of receiving the pension and its value in 1990 or before. I focus on girls since no effect are found for boys. There is little catch up growth occuring after 24 months. It therefore seems reasonable to assume that α must vary between 0 and 5%. The probability of receiving the maximum pension before 1990, conditionally on being eligible, is not well documented. I will assume that the program at least doubled its coverage from 1990 to 1993. The maximum amount versed each year are taken from table 1.

The first 2 columns in table 8 provide four possible point estimates of the value of receiving the pension, based on the coefficient estimated in column 2. The lowest estimate is 0.58 and the highest is 1.22. The average height for age of eligible girls is -1.23. Therefore, these results suggest that the pension helps them bridging at least half the gap with American girls their age. Columns 3 and 4 provide estimates of the coefficient of each rand received. The coefficients range from 0.14 to 0.29 per 100 rand. The lowest of this point estimates suggests that a transfer of one dollar a day would raise the height for age of the girls by 0.21 standard deviations.

It would be interesting to compare this effect with the effect of income from other sources. Unfortunately, we cannot estimate consistently the effect of a non-pension income on height for age. Non-pension income is not an exogenous variable. Moreover there is considerable measurement error in non-pension income. It is therefore not clear whether OLS estimates of the effect of pension income are biased upwards or downwards. To limit the extent of measurement error, I estimate by 2SLS a specification where I regress height for age on non-pension income, using eligibility as an instrument for pension income, and a series of indicator variables as

instruments for non-pension income.[19] I also restrict the sample to positive incomes and income smaller than 1900 rands (the 90% poorest households). Results are presented in column 5 of table 7. For boys, I find small and insignificant effect of non-pension income on height for age, which is consistent with the fact that we do not find strong evidence of an effect of pension income. For girls we find that both non-pension income and pension income have an effect. The point estimate of the effect of a rand of pension income is larger than the effect of a rand of non-pension income, but standard errors are too large to reject that they are equal.

At least two interpretations of this set of results are possible. First, if households are credit constrained, this can lead them to focus their investment on improving the health of boys, presumably because returns to the health of men (or boys) are higher (Garg and Morduch (1998) present evidence of this effect in an other African country, Ghana). If the pension relaxes the credit constraint, girls will then benefit from this more than boys. Another interpretation is that the tastes of the elderly persons are different from the tastes of prime-age household members, and that elderly persons have a preference for their granddaughters' health. In section 5, I examine whether the effect of the pension varies with the gender of the recipient. In the next section, I discuss the identification assumptions and propose additional evidence to address some potential problems.

5 Discussions of the identifying assumption and additional evidence

5.1 Endogenous household formation

5.1.1 Endogenous household formation as a source of bias

Until now, I have assumed that households' eligibility status was not affected by the program. However, changes is family composition might have occurred as a response to the program. Living arrangements under which children are living with their grand-parents (with our without their parents) were frequent before the pension program. But some living arrangements might

[19] The instruments are: head of the household is employed, head is self-employed head holds a regular wage job, a casual wage job, a job in agriculture, sector of head's job, head works for the Government or an NGO, head works for a private firm, head is paid monthly, fortnightly, weekly.

have been modified as a result of the pension. This could affect the validity of the identification strategy proposed in this paper. Assume for example that parents who care the most about the health of their children have sent them to live with their grand-mother after she started receiving the pension. This will imply that children who live with an eligible member come from families who care more about their health, and therefore they might be taller than their peers for reasons other the increase in income *per se*. If this effect is additive (for example if "good" families send *all* their children to live with their grand-mother after she starts receiving the pension), this does not invalidate the spirit of the strategy, which is to compare the differences between old and young children in eligible and non-eligible households. But if health-conscious parents send only their youngest children to live with their grand parents, younger children might be taller than their peers in eligible families for reasons other than the pension. Conversely, if health-conscious parents send only the oldest children to live with their grand-mother after she starts receiving the pension, this will results in downward biased estimate of the effects of the program.

5.1.2 Alternative identification strategy

To address this problem, I use an alternative instrument, which is designed to be correlated with the presence of an eligible member in the household, but not affected by household decisions. This instrument is a variable that indicates whether the child has at least one grand parent who is alive and eligible, or likely to be eligible. The survey instrument asks household members whether their parents are alive. This makes it possible to determine whether the child has at least one grand-parent alive, but not to establish whether the grand-parent is eligible (unless if he or she does not live in the household). However, if the mother and father are old enough their parents (if alive) are likely to be eligible. In practice the instrument is equal to 1 if there is an eligible person in the household or if one of the following is true: the mother (the father) of the child is older than 34 and her (his) mother is alive or the mother (the father) of the child is older than 32 and her (his) father is alive.[20] 46% of children who have an old grand parent

[20]I determined the cutoffs of 32 and 34 years by using the information on extended families in my sample. Women whose observed child is above 34 and men whose observed child is above 32 have a probability of 60% to be eligible for the pension. Results are not sensible to the choice of the cutoff. If a parent is not in the household, the survey does not indicate nor his age, nor whether his parents are alive. So some children may have an old grand parent alive not identified in my data.

alive (and 7.8% of those who do not) live with a pension recipient. Therefore this instrument is still strongly correlated with pension receipt. We can use it as an alternative instrument for pension receipt, and check whether results are consistent with those obtained using eligibility.

90% of children have at least a grand parent alive. The variation in the instrument comes mostly from parents' age. Children who have older parents may either be children who have many siblings or children whose parents started to have children older, which have opposite consequences on their welfare. In practice, the two sub-samples (where this instrument is 0 or 1) are more similar to each other than eligible and non eligible households are. This is apparent in the descriptive statistics presented in table 2 (columns 6 and 7). Two sets of characteristics are (not surprisingly) different: the child's parents are on average older and family size is on average larger when the child has an old grand parent alive. This increased similarity is reflected in their height in panel C. The difference in height for age between children is -0.04 (against -0.21 between eligible woman no eligible member). The difference in differences estimator will control for systematic differences in the two groups which is fixed across cohorts, and is not affected by endogenous family formation.

Descriptive statistics broken down by age (presented in table 3) have the expected pattern. Among girls born after January 1992, girls who have an old grand-parent alive are taller than other girls. The opposite is true for girls born before January 1992. For boys, we see a closing of the gap, but boys born after January 1992 are still smaller if they have an old grand-parent alive than otherwise. This suggests that using this instrument instead of eligibility directly should leave conclusions unchanged.

Figure 3-B shows the difference between the height in centimeters of children who have a grand parent alive and that of other children. The pattern is similar to that in figure 1-B: the difference is at first slightly negative, then the gap closes, and finally it becomes positive. Figure 3-A shows non-parametric regressions of height-for-age as a function of date of birth for children who have an old grand parent alive (straight line) and other children (broken line). Not surprisingly, the differences before the program are less marked in this case. However children born after January 1992 are definitely taller if they have a grand-parent alive and old.

Average derivatives of the difference between these estimated functions are presented in table 4 (columns 3 and 4). Here too, we find the same pattern as for eligibility, although the estimates are barely significant, even for girls: before January 1991, the average derivative is close to 0 for

all measures, for boys and for girls. After January 1991, the average derivative is more positive for girls, and about twice as small as the average derivative of the difference between eligible and non eligible (column 4) -which is expected since, this indicator is less tightly knit with the probability of receiving a pension. These numbers, however, are not significant at more than 80% confidence. For boys, we find a smaller, positive average derivative.

Estimates of a difference in differences specification similar to the reduced form equation 6, but where I use the indicator for whether the child as an old grand parent alive instead of eligibility status are presented in table 5 (columns 4 to 7). For girls, I find a positive (but insignificant) effect of having a grand parent alive, smaller than the estimated effect of eligibility. This is what we expected, since the probability of getting the pension is larger conditional on living with an eligible member than conditional on having an old grand parent alive. The effect of having a grand parent alive and old, uninteracted, is small and not significant. The two groups compared are now similar. I now find a positive difference in differences (even though it is not significant either) for boys as well. Taken together, the results for boys suggest that there might be an effect of the pension for boys, which is confounded by the fact that younger boys living with their grandmothers are systematically worse off than older boys. The fact that this is true for boys, but not for girls, is not particularly surprising: fostering practices are highly gender specific throughout the continent. Alternatively, boys could benefit from inter-household transfers. Jensen (1998) shows that inter-generational transfers responded to the program. Transfers from children to parents were reduced when parents got eligible. Some non co-resident grandparents may also have started remitting to their children when they became eligible. Money transferred is not directly spent by the pension recipient, but by the recipient of his transfer. As we will see below, the identity of the income recipient matter for the disposition of the income. It is therefore quite possible that money received (or not sent) by the children's parents was spent toward the boys rather than toward the girls.

We can then compute 2SLS estimates of the effect of pension receipt, (equation 7) using the interaction between the indicator for having an old grand-parent alive and being born after January 1992 as an instrument. The first stages are shown in table 6 (columns 2). There is still a strong relationship between the receipt of the pension and this instrument. The coefficient is 0.54 with T statistics above 14. The 2SLS estimates are shown in table 7 (column 4). For girls, the estimates are less precise than when using eligibility as an instrument, but the point

estimate of the effect of pension using this instrument is very close from the corresponding estimate in column 2. (0.58 and 0.59, respectively). This result indicates that endogenous family composition does not bias the estimates of the effects of eligibility for girls. For boys, I now find a positive point estimate, almost as large as that of girls (0.50).

Using this alternative instrument leads to a reduced precision, but using it does not change the conclusions for girls, and confirms that the pension had an effect on their nutrition. For boys, it now indicates that there may be some effect of the pension, although the estimates are too imprecise to be definitive.

5.2 Omitted variables. Additional evidence on weight for height

Another potential problem for the interpretation of differences in differences is that unobserved factors correlated with receiving a pension might have different effects at different ages. For example, unobserved quality of the family might be a stronger determinant of height for age for older children than for younger children. Alternatively, another program, which targets the same type of households, could have started at the same time as the pension program. This is a period of great change in South Africa, and a large portion of these changes may have affected households who have an eligible member, a poorer and more rural population than the rest of South Africa, disproportionately.

Note, however, that the non-parametric regressions of height for age as a function of age in eligible and non eligible households actually *cross*. Younger girls are taller in eligible households and the reverse is true for older girls (see figures 1 and 2 and table 5). This pattern cannot be explained by an unobserved variable affecting younger children differently.

This pattern could also be caused by another program that affected young children in eligible families disproportionately, an immunization program for example. However, such a omitted variable would also be reflected in weight for height. Whereas height for age reflects mostly nutrition since birth, weight for height reflects nutrition and illnesses within the past few weeks. This suggests that if the differences in differences are due to improved nutrition and not to differential effects of unobserved family characteristics (or to endogenous family formation affecting only the youngest children), there should be no differences in differences in weight for height. The nutrition of all children, not only the youngest ones should have improved as a result of the pension. The difference between the weight for height of children in eligible and non

eligible families should therefore not vary with age. In this subsection, I first present evidence suggesting that weight for height of girls (but not that of boys) is higher for children living in eligible families (table 9), and then that the effect on older and younger children is similar.

In table 9, I present direct evidence on weight for height, which reinforces the finding that the program improved the nutrition of girls but not of boys. In column 1, I regress the weight for height Z-scores on eligibility status and a dummy for whether the household has an eligible member. For boys, the coefficient is negative, while it is positive for girls. Column 2 shows that the effect of having a elderly household member is not positive for girls if the elderly person is not eligible. Having a member within 5 years of eligibility (but not eligible) in the household has a negative effect on the weight for height of girls and no effect on that of boys. Having a member within 5 to 10 years of eligibility has no effect on the nutrition of either boys or girls. The fact that there is no positive effect of having a member close to eligibility fits with our earlier result that the difference in height for age between eligible and non eligible households only narrowed for children born during the expansion. This suggests that households cannot borrow against expected pension funds. In column 3, I regress weight for height on eligibility, controlling for whether there is at least one household member above 50.[21] The coefficient on the weight for height of boys is still insignificant, while that of girls doubles. The next three columns show that these results are insensitive to additional control variables. These results, which do not exploit the age of the children but only that of the grandparent, correspond exactly to the results on height for age: only the weight for height of girls is affected by the program.[22]

Next, I examine whether the effect on weight for height is larger for younger children. Panel C in table 3 shows means of weight for height in various sub-samples. Boys and girls of all ages are heavier if they live with an eligible woman than if they live with no eligible member. The difference is larger for girls than for boys which is consistent with previous results.

The average derivatives of the difference between weight for height in eligible and non-eligible families as a function of date of birth are presented in table 4 (panel C). They are very small and insignificant, suggesting that there is no significant pattern of the eligible children doing

[21] This is a specification very similar to that of Bertrand et al. (1999) or Case and Deaton (1998).

[22] There is some concern about the quality of the weight data. This significant (and expected) result suggests that it is not pure noise. I checked whether there was suspect bunching in the data, which would suggest guessing. There is definitely bunching at entire values of weight, which indicates that the measurements were not too precise, but not around multiples of 5, which would be an indication of bunching.

better over time.

In table 10, I present estimates of difference in differences reduced form specifications similar to equation 6, but where weight for height is the dependent variable. In contrast to previous results, there is no significant difference in differences in this table. For girls and boys the coefficients of the interaction are very small and insignificant. The effect of eligibility is smaller and not significant for boys.

Another potential control experiment is height for age in groups whose pension benefits did not increase during the 1990s. This includes all racial groups but Africans. Pension benefits for Whites actually declined somewhat in the 1990s (although not as fast as they did in the 1980s). There are a few non-Africans in the sample, which precludes a very conclusive analysis. However, I ran the differences in differences estimate in the sample of non-African children, and I actually found negative and insignificant differences in differences. This result, in addition to the weight for height results, reassures that the results found for Blacks are not spurious.

It is possible, in principle, that another program could have affected the nutrition of a group of children with similar backgrounds in the same period. Three pieces of evidence contradict this view. First, I obtain similar results when I use the indication for a grandparent alive and elderly as when I use the indication for eligibility (if anything, larger for boys), despite the fact that the control and treatment groups are similar in this case (there is no significant negative effect of having an elderly grandparent alive among older children). Second, such a program could potentially target three generation households (because they are poorer), but not in particular households where a grandparent is eligible. However, the evidence on weight for height shows that it increased only for girls, and only in households where a member was eligible, not in those where a member was close to eligibility. Moreover, when I control for the presence of a member above 50, interacted with year of birth in the height for age regression, the estimate increases. This indicates that the positive difference in difference is likely to be due to the pension, and not to another ignored program. Finally, evidence will be provided below that a pension received by a grandfather has no effect on child growth. Yet, as shown in table 2, the characteristics of the households with a man or a woman eligible are very similar. It is, therefore, reasonable to expect that they would be targeted by the same programs.

6 Further results. Differences by Gender

This section provides a more detailed analysis of further results and their implications on the effectiveness of cash transfers on child outcomes. In particular, it investigates whether the transfers have bigger effects in some groups than in others, and when they are received by women rather than by men.

6.1 Effects on different groups

The effect of the pension is expected to vary across group. Table A1 present the differences in differences estimates of the effect of eligibility of the pension) in different sub-samples (with and without controlling for the presence of a grandparent).

One first source of variation is the level of wealth of the household. The effect should be smaller in richer households. The estimates, presented in table A1 are indeed larger among poor households. To avoid the problems caused by the fact that non-pension income is not pre-determined with respect to the receipt of the pension[23], in column 3 and 4, I show the estimated effects among children whose mother's education is less than the median (5 years) those whose mother's education is above median. For girls, these two ways of splitting the sample give similar results: the point estimate of effect of the pension is larger in the sample of poorer (or less educated) households than in the sample of richer households and in the sample of less educated mothers than in the sample of more educated mothers.[24] For boys, there little or no effect in all sub-samples, except perhaps among educated mothers.

The effect should be smaller when the pension is worth less to the children. It is worth less to each child when more children have to share the pension. It is also worth less when the pension recipient is very old, and thus is expected to die sooner. Columns 5 and 6 show the results obtained by splitting the sample according to whether the household has fewer or more than the median number of household members aged 16 or younger (which is four). The effect is bigger when the number of children is smaller. Columns 9 and 10 interact eligibility with the age of the eligible member. There is no effect of eligibility when the eligible member is older than 70 (50.4% of eligible member are above 70).

[23]Bertrand et al. (1999) argue that the receipt of the pension is associated with a lower labor supply by prime-aged adults in the household.

[24]The standard errors are large, however, therefore these differences are not significant.

Finally, when there are two eligible members they receive the equivalent of two full pensions. Column 7 and 8 use a separate indicator to indicate whether there are one or two eligible members in the household. For boys, two eligible members have no more positive effect than one. For girls, the point estimate of two eligible member is slightly larger than the point estimate of one eligible member, but very close. A single eligible member is very often a woman. This suggest that the role of an eligible man, as the second eligible member, is not very important. The next subsection further investigates this issue.

6.2 Importance of recipient's gender

Three out of four pension recipients are women (they live longer and they are eligible earlier). The pension is therefore a transfer program biased in favor of women. It has been argued (Lund (1993)) that it is a desirable feature, since income in the hands of women tend to be more strongly associated to "desirable" outcomes (heath, education, etc...) than income in the hand of men. However, this is based on evidence from elsewhere, which is moreover subject to the caveats mentioned above. The reform in the pension program is a unique occasion to examine whether pension income has a different impact in the hands of women than in the hand of men. The comparison is not affected by the two problems (endogeneity of income and functioning of the marriage market) that plague most previous efforts to establish similar results in the literature. Does it matter that pension income is received by a woman or a man?

6.2.1 Descriptive statistics and Non-parametric results

The descriptive statistics in table 3 indicate that the gender of the eligible person matters . Children of both sexes are on average smaller if they live with an eligible man than with an eligible woman. Moreover, boys born before January 1992 have the same average height in both cases, whereas younger boys living with an eligible man are much smaller than other boys. For girls, the pattern is even more striking. Older girls do better when they live with an eligible man than if they are living with an eligible woman, while the opposite is true for younger girls.

Non-parametric regressions illustrate these differences. In figure 5-A, the plain line is the difference between the height of children in households where there is an eligible man and that of children in households where there is no eligible member. The dotted line is the difference between the height of children in households where there is an eligible woman and that of

children in households where there is no eligible member. [25] For eligible women, the difference starts negative, and decreases in absolute from January 1991 on before becoming positive after January 1992. For men, the difference is less negative before 1991, and becomes less positive afterwards. This figure suggests a bigger impact when the pension is received by women than when it is received by men.[26]

6.2.2 Differences in differences and 2SLS estimates

I extend the difference in differences formulation to take into account the gender of the eligible individual.

$$h_{ifk} = \pi_w 1_{(k=1)} * TW_f + \pi_m 1_{(k=1)} * TM_f + \beta_w TW_f + \beta_m TM_f + \sum_{j=1}^{3} \gamma_j 1_{(j=k)} + \epsilon_{ifk}, \qquad (9)$$

where TW_f is equal to 1 if there is an eligible woman in the household and 0 otherwise and TM_f is equal to 1 if there is an eligible man in the household and 0 otherwise.[27] A similar formulation is estimated with TW_f being equal to 1 if the child as a old grand mother alive and 0 otherwise and TM_f being equal to 1 if the child has a old grand-father alive (this instrument is defined exactly as before). Finally, the effect of male and female pension receipt can be estimated by using the interactions between eligibility by gender and dummy for being young for the interactions of pension receipt by gender and the dummy for being young.

I present the estimates of equation 9 in table 11. For boys as well as girls, the difference in differences are positive for woman's eligibility or but negative (and not significant) for men's eligibility. The positive effects of woman's eligibility is larger on the height of girls than on that of boys, and it is significant only for girls.

The results on weight for height presented in table 9 are, once again, consistent with the results on height for age. In the last column of this table, I present the results of a regression of the child weight for height on an indicator for woman eligible, man eligible, woman above 50

[25] In some households where there is an eligible man, there is also an eligible woman. In this version of the graphs these households are included in both samples. This tends to attenuate any difference between the two regressions.

[26] Using height for age instead of height or the alternative instrument lead to similar conclusions.

[27] I have omitted family background variables in this notation, but I introduce them when I estimate the equation.

and man above 50. Girls are significantly heavier if they live with an eligible woman, controlling for the presence of a woman above 50 (the point estimate is 0.50, with a a standard error of 0.17). The effect of living with an eligible man is close to 0 and insignificant. For boys, both the effect of woman's and man's eligibility are negative and insignificant.

A potential explanation for this difference is that grandfathers are more likely to transfer money to their non co-resident children, which would attenuate the difference between eligible and non eligible households. However, using the indicator of whether the child has a grand-mother or a grand-father alive and old leads to the same conclusion. The effect of having a old grand mother alive is positive for both boys and girls, larger for girls than for boys, and significant only for girls. There is a small and insignificant positive effect on boys of having a male grand-parent alive, and a small and insignificant negative effect on girls.

Men are more likely to have worked than women. They could be less likely to receive the pension if many of them had employer's pension. In practice, however, the difference is not large. The first stage relationship shows that the probability to receive a pension conditional on begin eligible is 0.62 for a man and 0.74 for a woman. The 2SLS results in table 11 (column 4) establish clearly that a pension received by a woman is effective (especially on girls), while a pension received by a man is not. The point estimates suggest that pension received by women lead to an increase in the height-for-age of boys by 0.26 (this is not significant) and that to that of girls by 0.69. In contrast, the estimates of the effect of a pension received by a man is negative and insignificant.

6.2.3 Interpretation

These results provide an example of the differential impact of women's and men's income on child health which is not subject to the traditional caveats in this literature (measurement error, endogeneity of income and correlation between woman non-labor income and unobserved man's characteristics due to marriage). However, there two interpretations are still possible. The first interpretation is that the household is a collective, not a unitary, entity, and that the same resources are spent differently when they are received by a woman and when they are received by a man. Another interpretation, however, could be that, in term of permanent income, a rand of pension income received by a man represents much less than a rand of pension income received by a woman, because men receive the pension for a shorter time. If household are credit

constrained or have a very high discount rate, this should not lead to different effects of man's and woman's pensions. But we have seen that the age of the eligible member affects the effect of the pension.

To check whether the health status of the eligible member affects the results, I first directly control for the health status (a dummy for whether the individual has been sick in the period preceding the survey) of the eligible household member, interacted with cohort dummies. This specification is reported in column 2 of table 11. This leaves estimates essentially unchanged.

In addition, to help discriminate between these two interpretations, it is useful to look at the disposition of men's an women's pension income. If the household is a unitary entity, and if a man's pension income is not spent on child health because it is akin to transitory income, then we should see that the propensity to save out of men's income is much larger than the propensity to spend out of women's income (and non-pension income). I therefore estimate the following equation:

$$S_f = \alpha_w y_{fw} + \alpha_m y_{fm} + \alpha z_f + X_f \beta + \epsilon_f, \tag{10}$$

where S_f stands for the total savings of households (defined as total income minus expenditures), y_{fw} is pension income received by women, y_{mw} is pension income received by men, z_f is non pension income, and X_f is a set of control variables. This specification extends Case and Deaton (1998) formulation to take into account differences in the disposition of income received by men and women. The emphasis here is on the comparison between α_w and α_m. This equation is estimated by OLS, and 2SLS. The instruments in the 2SLS equations (y_{fm}, y_{fw} and z_f are instrumented) are the indicators for the presence of an eligible man and an eligible woman, and the instruments used to correct for measurement errors in non-pension income, which were discussed earlier.

Results are presented in table 12 (column 1 and 2). The point estimates suggest that propensity to save out of a man's pension income is actually *lower* than the propensity to save out of a woman's non-pension income (although this difference is not significant). This result indicates that the differences in the effects of woman's and man's pension income on child height is not likely to be due to their different life cycle properties. In combination with the results in the previous subsection, this result therefore suggests that the disposition of income is influenced

by the gender of the recipient.

6.3 Importance of the intermediate generation

Finally, it is interesting to study the role of the intermediate generation. If the grand-parent directly spent the pension on what they want to buy (for example, children food), we could still see an impact of the intermediate generation, if the grand-parent had a preference for their daughter's children versus their son's children. It could also be that the pension recipient gives at least part of the money to her son or daughter. The preference of the intermediate generation will then be directly taken into account. If some cases, it is possible to determinate whether the eligible member is the parent of the child's mother or the child's father. In 34% of the eligible families, the eligible member is the mother's parent. In 45% it is the father's parent. In the remaining cases, it is not determinate. Column 5 in table 11 presents the results of estimating an equation similar to equation 9, but allowing the effect to vary according to the gender of the intermediate generation. The results are striking: For girls, we do find a strong effect when it is the mother's parent who is eligible, and a much smaller effect when it is the father's parent. When the sample is restricted to not include eligible males, the contrast is even larger.

The effect of the pension is gender specific at all levels. This can be compared with the results in Bertrand et al. (1999) on the effects of the pension on prime-age members labor supply. They find that the pension reduces the labor supply of prime-age males only, and that this effect is bigger when it the pension is received by a woman than when it is received by a man. The findings in these two studies, taken together, are consistent with a model of the family where males use the money they receive for their private consumption, while females dispose of the money in a more altruistic way, and, in particular, give it to their children.

7 Conclusion

The extension of the Old Age Pension program in South Africa has led to an improvement in the health and nutrition of children, especially girls. This is reflected in the height for age of the youngest children affected by the program expansion. This effect is entirely due to pensions received by women. I estimate that a pension received by women improved the height for age Z-scores of girls by at least 0.69 of a standard deviation and that of boys by 0.26 of a

standard deviation. South African children are on average -1.28 standard deviations smaller than American children, so it is a large increase. This is because the pension benefits are generous, but the effect of each rand of the pension is also important. In contrast, pension received by men have no effect on the height of children.

These findings contribute to answering the questions posed in the introduction. First, the fact that pensions received by women led to a sizeable increase in the height of girls shows that cash transfers can have an important effect on child health. Second, the fact that the pension had no effect on child height if it was received by a man or by the paternal grandmother, suggests that concerns about the excessive liquidity of cash transfers (relative to transfers in kind) are justified. The allocation of a cash transfer within the household seems to depend very strongly on the recipient of the grant. Even if we could reliably measure the average effect of household resources, this would not necessarily indicate the effect of a transfer. The marginal impact of the transfer may be very different from the average impact of income. This has implications for the design of cash transfers. In South Africa, the program is naturally biased toward women, both because men can claim the pension only after age 65, and because women tend to live longer. Without this feature, the program would not affect the nutrition of young children as much. The distinction between men and women is not in accordance with the South African constitution, and there is some pressure to remove it. The effectiveness of the pension program as a tool to transfer resources to young children would suggest moving in the opposite direction. However, other methods of investment in child human capital that we are not measuring here (education for example) could be affected by pensions received by men, so these implications needs to be taken carefully.

Finally, targeting elderly persons might be a way to reach a fair number of poor children in settings where a means test is not practical. Moreover, with the increasing prevalence of AIDS, targeting grandparents (not necessarily grandparents who are older than 60 or 65) might be the most effective way to reach children. Future work needs to compare the effectiveness of the pension and that of a child grant (which the South African government is trying to put in place) in reaching poor children and, once they are reached, in improving their health.

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Table 1: Maximum payable pension
(annual amount in constant 1990 Rand)

	Pension amount	
	Whites	Blacks
1987	3871	2077
1988	3431	1842
1989	3439	2054
1990	3308	2096
1991	3266	2444
1992	3152	2677
1993	3081	3081

Source: Van der Berg (1994)

Table 2: Descriptive statistics

	Receipt of pension		Eligibility for pension			Grandparent alive and old	
	Yes	No	Woman	Man	None	Yes	No
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: HOUSEHOLD CHARACTERISTICS							
Father's age	36.5	37.0	35.5	40.9	36.8	38.9	33.1
	(0.87)	(0.38)	(0.79)	(2.10)	(0.37)	(0.43)	(0.59)
Mother's age	28.0	29.6	28.2	27.5	29.5	31.4	26.5
	(0.34)	(0.20)	(0.38)	(0.58)	(0.20)	(0.24)	(0.24)
Mother's education	5.67	5.19	5.70	5.78	5.17	5.08	5.63
	(0.15)	(0.085)	(0.16)	(0.24)	(0.086)	(0.11)	(0.11)
Father's education	4.97	4.52	5.07	4.20	4.54	4.62	4.76
	(0.26)	(0.11)	(0.27)	(0.46)	(0.11)	(0.14)	(0.17)
Rural residence	0.75	0.67	0.75	0.83	0.67	0.70	0.67
	(0.018)	(0.012)	(0.018)	(0.028)	(0.012)	(0.014)	(0.015)
Grandparent in	0.93	0.42	0.95	0.89	0.42	0.56	0.56
household	(0.0099)	(0.012)	(0.0081)	(0.021)	(0.012)	(0.014)	(0.015)
Father is absent	0.64	0.42	0.67	0.66	0.41	0.41	0.55
	(0.019)	(0.012)	(0.020)	(0.033)	(0.012)	(0.014)	(0.015)
Mother is absent	0.18	0.08	0.18	0.14	0.08	0.10	0.11
	(0.014)	(0.0059)	(0.016)	(0.023)	(0.0059)	(0.0082)	(0.0089)
Household size	10.0	7.8	10.5	10.5	7.6	8.7	7.7
	(0.17)	(0.093)	(0.21)	-0.3	(0.086)	(0.13)	(0.11)
PANEL B: PENSION RECEIPT							
Male receives	0.32	0	0.17	0.68	0.03	0.14	0.04
pension	(0.019)	0	(0.016)	(0.034)	(0.0041)	(0.011)	(0.0059)
Female receives	0.83	0	0.79	0.47	0.04	0.40	0.05
pension	(0.015)	0	(0.018)	(0.037)	(0.0050)	(0.014)	(0.0068)
Non-pension income	665	927	723	637	908	813	885
	(33)	(22)	(36)	(51)	(22)	(24)	(29)
Pension income	383	0	325	389	23	182	27
	(6.6)	0	(9.6)	(20)	(2.2)	(6.7)	(3.2)
Per-capita income	123	148	121	123	149	146	143
	(3.9)	(3.9)	(4.5)	(7.3)	(3.9)	(4.3)	(4.7)
PANEL C: ANTHROPOMETRIC DATA							
Height for age Z-score	-1.42	-1.19	-1.38	-1.46	-1.21	-1.28	-1.23
	(0.067)	(0.036)	(0.072)	(0.13)	(0.036)	(0.046)	(0.045)
Weight for age Z-score	0.17	0.16	0.28	0.12	0.15	0.12	0.21
	(0.04)	(0.04)	(0.08)	(0.15)	(0.04)	(0.05)	(0.05)
Number of observations	923	2407	816	286	2380	1768	1562

Notes: Standard errors in parentheses.

For household level variables, the table present household averages weighted by the number of children in each household (multiplied by survey weights). For anthropometric data, the table presents individual level averages.

Table 3: Means of height for age by age groups

	Receipt of pension		Eligibility for pension			Grandparent alive and old	
	Yes	No	Woman	Man	None	Yes	No
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Boys born 01/92 or later	-1.30	-1.08	-1.20	-1.8	-1.09	-1.11	-1.16
	(0.17)	(0.10)	(0.19)	(0.37)	(0.10)	(0.13)	(0.12)
Boys born before 01/92	-1.59	-1.34	-1.62	-1.63	-1.32	-1.49	-1.32
	(0.10)	(0.053)	(0.11)	(0.17)	(0.053)	(0.068)	(0.064)
Girls born 01/92 or later	-0.85	-0.88	-0.62	-0.85	-0.94	-0.75	-0.99
	(0.21)	(0.11)	(0.22)	(0.41)	(0.11)	(0.14)	(0.13)
Girls born before 01/92	-1.48	-1.26	-1.47	-1.29	-1.26	-1.34	-1.30
	(0.11)	(0.059)	(0.12)	(0.045)	(0.059)	(0.073)	(0.076)

Note: Standard errors in parentheses.

Table 4:
Average derivative of the difference $g^{Treatment}$(date of birth)-$g^{control}$(date of birth)

"Treatment" / "Control"	Eligible / Non-eligible		Grandparent alive and old / No grandparent alive and old	
	06/88-12/90	01/91-06/93	06/88-12/90	01/91-06/93
	(1)	(2)	(3)	(4)
PANEL A: HEIGHT IN CENTIMETERS				
A.1 Boys				
Average derivative	-0.044	0.037	-0.016	0.016
	(0.044)	(0.035)	(0.037)	(0.032)
A2. Girls				
Average derivative	0.009	0.093	0.010	0.047
	(0.045)	(0.045)	(0.044)	(0.036)
PANEL B: HEIGHT FOR AGE Z-SCORES				
B.1 Boys				
Average derivative	-0.013	0.013	-0.007	0.004
	(0.010)	(0.011)	(0.009)	(0.010)
B2. Girls				
Average derivative	0.001	0.026	0.001	0.011
	(0.019)	(0.014)	(0.008)	(0.012)
PANEL C: WEIGHT FOR HEIGHT Z-SCORES				
C.1 Boys				
Average derivative	-0.007	0.011	0.000	0.018
	(0.011)	(0.014)	(0.0041)	(0.011)
C2. Girls				
Average derivative	-0.007	-0.012	-0.008	0.007
	(0.011)	(0.014)	(0.010)	(0.012)

Note: Standard errors (robust to correlation of residuals within households and to heteroscedasticity) in parentheses.

Table 5: OLS regressions: Impact of eligibility status on height for age

	Status=Eligible for pension			Status=Grandparent alive and old		
	(1)	(2)	(3)	(4)	(5)	(6)
PANEL A: Boys (N=1272)						
Status*born after 01/92	0.076	0.046	0.08	0.27	0.27	0.31
	(0.27)	(0.27)	(0.31)	(0.22)	(0.25)	(0.24)
Status	-0.18	-0.18	-0.16	-0.16	-0.27	-0.17
	(0.14)	(0.14)	(0.26)	(0.12)	(0.13)	(0.14)
PANEL B: Girls (N=1234)						
Status*born after 01/92	0.44	0.44	0.56	0.26	0.28	0.24
	(0.27)	(0.27)	(0.34)	(0.23)	(0.28)	(0.26)
Status	-0.2	-0.16	-0.089	-0.03	-0.096	-0.083
	(0.15)	(0.15)	(0.19)	(0.13)	(0.14)	(0.14)
Covariates:						
Year of birth dummies	Yes	Yes	Yes	Yes	Yes	Yes
Family background variables *Year of birth dummies	No	Yes	Yes	No	Yes	Yes
Member's age variables *Year of birth dummies	No	No	Yes	No	No	Yes

Notes: Standard errors (robust to correlation of residuals within households and heteroscedasticity) in parentheses.
Family background variables: father's age and education, mother's age and education and rural or metro residence.
Member age variables: family size, number of members aged 0 to 5, 6 to 15, 16 to 24, 25 to 49, 50 and above.

Table 6: First stage regressions

	Dependent variable					
	Household receives pension		Pension income (in rands)		Woman receives pension	Man receives pension
	(1)	(2)	(3)	(4)	(5)	(6)
PANEL A: Boys						
Household eligible	0.74		288			
	(0.039)		(22)			
Grandparent alive and old		0.54		214		
		(0.039)		(20)		
Woman eligible					0.74	0.044
					(0.042)	(0.035)
Man eligible					-0.049	0.62
					(0.055)	(0.065)
PANEL B: Girls						
Household eligible	0.75		313			
	(0.041)		(23)			
Grandparent alive and old		0.54		226		
		(0.045)		(22)		
Woman eligible					0.69	0.033
					(0.052)	(0.033)
Man eligible					0.12	0.66
					(0.096)	(0.091)

All variables are interacted with a dummy for being born 01/92 or after.A Full set of family background covariates is included in all regressions. First stages are shown in the sample which excludes children born 01/91-12/91.

Table 7: OLS and 2SLS estimates
Effects of pension receipt on height for age

	Effect of receiving a pension				Effect of pension income (in rands, divided by 100)			
	OLS	2SLS			2SLS	2SLS		
Instruments:		Eligible *born afer 01/92		Grandparent alive and old *born afer 01/92	See note below the table	Eligible *born afer 01/92		Grandparent alive and old *born afer 01/92
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
PANEL A: Boys (N=1272)								
Receives pension or pension income *born after 01/92		0.060 (0.39)	0.14 (0.29)	0.50 (0.43)		0.015 (0.10)	0.036 (0.11)	0.13 (0.11)
Receives pension	-0.17 (0.12)							
Non-pension income					-0.0018 (0.052)			
PANEL B: Girls (N=1234)								
Receives pension or pension income *born after 01/92		0.58 (0.37)	0.82 (0.47)	0.59 (0.51)		0.14 (0.088)	0.20 (0.011)	0.14 (0.12)
Receives pension	-0.088 (0.13)							
Non-pension income					0.060 (0.028)			
Covariates:								
Eligible for pension	No	Yes	Yes	No	No	Yes	Yes	No
Grand-parent alive and old	No	No	No	Yes	No	No	No	Yes

Notes: Standard errors (robust to correlation of residuals within households and heteroscedasticity) in parentheses.
Instruments in column 5 are: dummies for: head is employed, head holds a regular job, a casual wage job, a job an agricultue, sector of the job employer's type (central or local government, private firm, other), pay type (weekly, fortnightly, monthly), dummy for eligibility for pension.
Equations estimated in a sample which excludes children born between 01/91 and 12/91. Household with non-pension income above 1900 rands monthly are excluded from column 5. All equations include year of birth dummies, family background variables, and family background variables interacted with year of birth dummies. Columns 4 and 8 include in addition household member age variables and household members age variables interacted with year dummies.

Table 8: Alternative computation of the effect of pension receipt (girls only)

		Effect of pension receipt		Coefficient of pension income	
		Catch-up parameter		Catch-up parameter	
		a=0	a=0.05	a=0	a=0.05
		(1)	(2)	(3)	(4)
Pre-1990 probability of receiving the old age pension among eligible	0	0.58	0.61	0.14	0.15
	0.35	1.16	1.22	0.28	0.29

Note: The estimate is taken from column 2 in table 7.
Pension income is expressed in rands, divided by 100.

Table 9: Effect of eligibility on weight for height: Cross sectional evidence

	Dependent Variable: Weight for height					
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: BOYS						
Eligible member	-0.028	-0.054	0.041	0.02	-0.011	
	(0.13)	(0.097)	(0.17)	(0.16)	(0.17)	
Eligible man						-0.023
						(0.26)
Eligible Woman						-0.17
						(0.19)
At least one member is:						
Within 5 years of eligility		0.085				
		(0.13)				
Within 10 years of eligibility		0.136				
		(0.12)				
Over 50			0.099	-0.16	-0.29	
			(0.16)	(0.16)	(0.20)	
At least one man over 50						0.11
						(0.19)
At least one woman over 50						-0.52
						(0.26)
PANEL C: GIRLS						
Eligible	0.22	0.20	0.40	0.40	0.37	
	(0.12)	(0.13)	(0.15)	(0.15)	(0.16)	
Eligible man						0.06
						(0.26)
Eligible Woman						0.50
						(0.17)
At least one member:						
Within 5 years of eligility	-0.34					
	(0.16)					
Within 10 years of eligibility	-0.08					
	(0.12)					
Over 50			-0.25	-0.25	-0.26	
			(0.13)	(0.14)	(0.17)	
At least one man over 50						0.09
						(0.15)
At least one woman over 50						-0.31
						(0.19)
Year of birth bummies	Yes	Yes	Yes	Yes	Yes	Yes
Family background variables	No	No	No	Yes	Yes	Yes
Household member's age	No	No	No	No	Yes	Yes

Notes: Standard errors (robust to correlation of residuals within households and heteroscedasticity) in parentheses.
Family background variables: father's age and education, mother's age and education and rural or metro residence.
Member age variables: family size, number of members aged 0 to 5, 6 to 15, 16 to 24, 24 to 49, 50 and above.

Table 10: Control experiment.

Differences in differences using weight for height as the outcome

	Status=Eligible for pension			Status=Grandparent alive and old		
	(1)	(2)	(3)	(4)	(5)	(6)
PANEL A: Boys (N=1240)						
Status*born after 01/92	0.12	0.14	-0.0078	-0.068	0.0087	-0.095
	(0.29)	(0.29)	(0.35)	(0.23)	(0.28)	(0.28)
Status	0.079	0.018	0.074	0.088	0.07	0.13
	(0.16)	(0.15)	(0.19)	(0.12)	(0.14)	(0.15)
PANEL B: Girls (N=1231)						
Status*born after 01/92	-0.010	-0.046	-0.01	-0.014	-0.052	-0.091
	(0.28)	(0.28)	(0.30)	(0.24)	(0.25)	(0.26)
Status	0.25	0.23	0.27	0.087	0.13	0.096
	(0.13)	(0.14)	(0.18)	(0.12)	(0.13)	(0.14)
Covariates:						
Year of birth dummies	Yes	Yes	Yes	Yes	Yes	Yes
Family background variables *Year of birth dummies	No	Yes	Yes	No	Yes	Yes
Member's age variables *Year of birth dummies	No	No	Yes	No	No	Yes

Notes: Standard errors (robust to correlation of residuals within households and heteroscedasticity) in parentheses.
Family background variables: father's age and education, mother's age and education and rural or metro residence.
Member age variables: family size, number of members aged 0 to 5, 6 to 15, 15 to 24, 25 to 49, 50 and above.

Table 11: OLS regressions
Effects of women's and men's eligibility on height for age

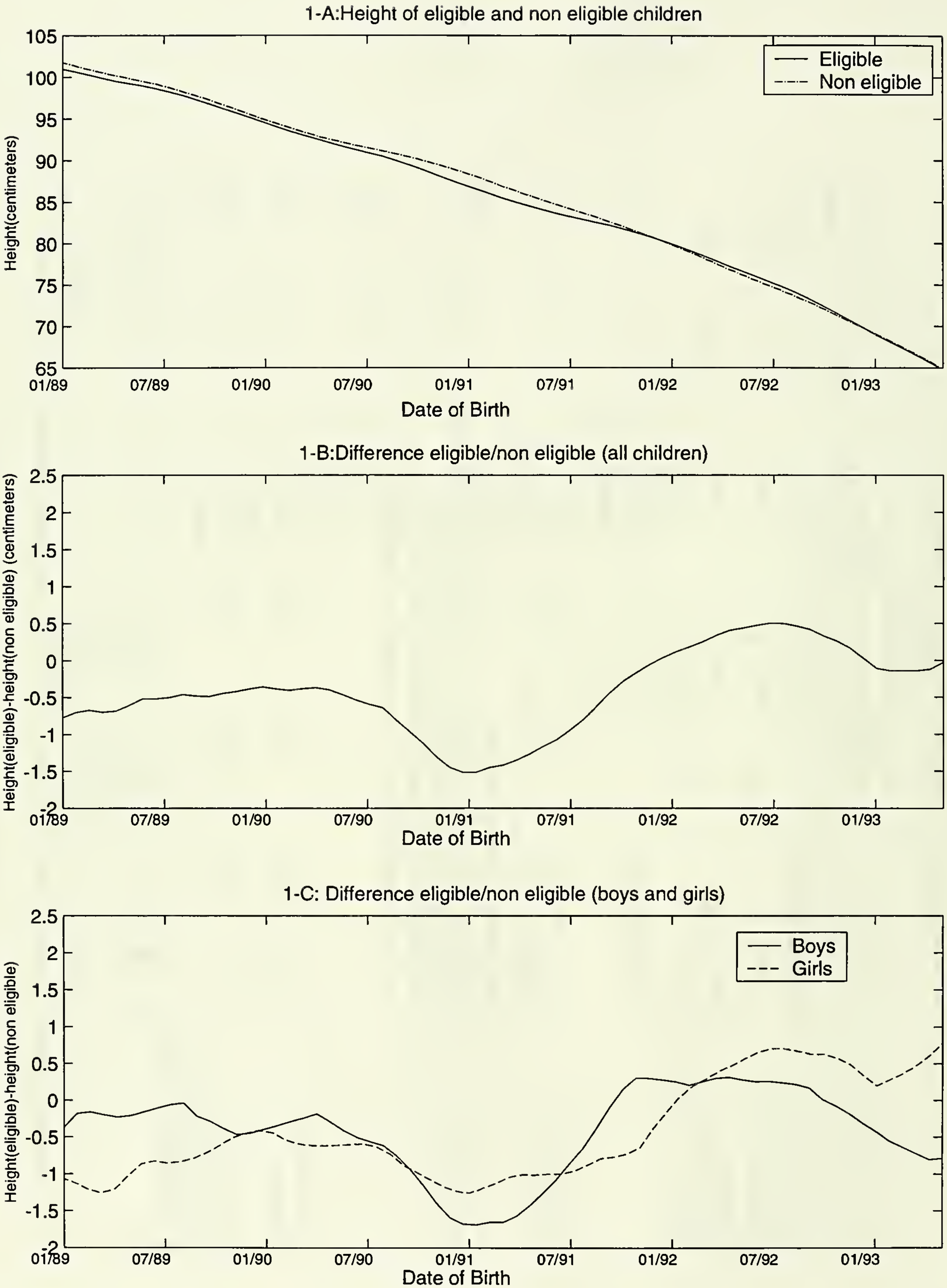
	Treatment						
	Eligibility		Grandparent alive and old	Receives pension		Eligibility	
	OLS		OLS	2SLS		OLS	
	(1)	(2)	(3)	(4)		(5)	(6)
PANEL A: Boys (N=1272)							
woman treated	0.16	0.14	0.11	0.26	Mother's parent	-0.018	0.35
*born after 01/92	(0.29)	(0.29)	(0.24)	(0.42)	eligible	(0.47)	(0.62)
					*born after 01/92		
man treated	-0.37	-0.45	0.21	-0.59	Father's parent	0.059	0.19
*born after 01/92	(0.35)	(0.45)	(0.26)	(0.67)	eligible	(0.35)	(0.37)
					*born after 01/92		
					Cannot tell which	-0.15	-0.24
					parent is eligible	(0.47)	(0.54)
					*born after 01/92		
PANEL C: Girls (N=1234)							
woman treated	0.47	0.48	0.48	0.69	Mother's parent	0.77	1.21
*born after 01/92	(0.29)	(0.31)	(0.24)	(0.43)	eligible	(0.58)	(0.87)
					*born after 01/92		
man treated	0.041	-0.09	-0.14	-0.089	Father's parent	0.22	0.15
*born after 01/92	(0.51)	(0.56)	(0.33)	(0.44)	eligible	(0.31)	(0.35)
					*born after 01/92		
					Cannot tell which	0.15	0.27
					eligible	(0.49)	(0.58)
					*born after 01/92		
Sample:							
Excludes eligible males	No	No	No	No		No	Yes
Covariates:							
Eligible person sick	No	Yes	No	No		No	No
*Year of birth dummies							

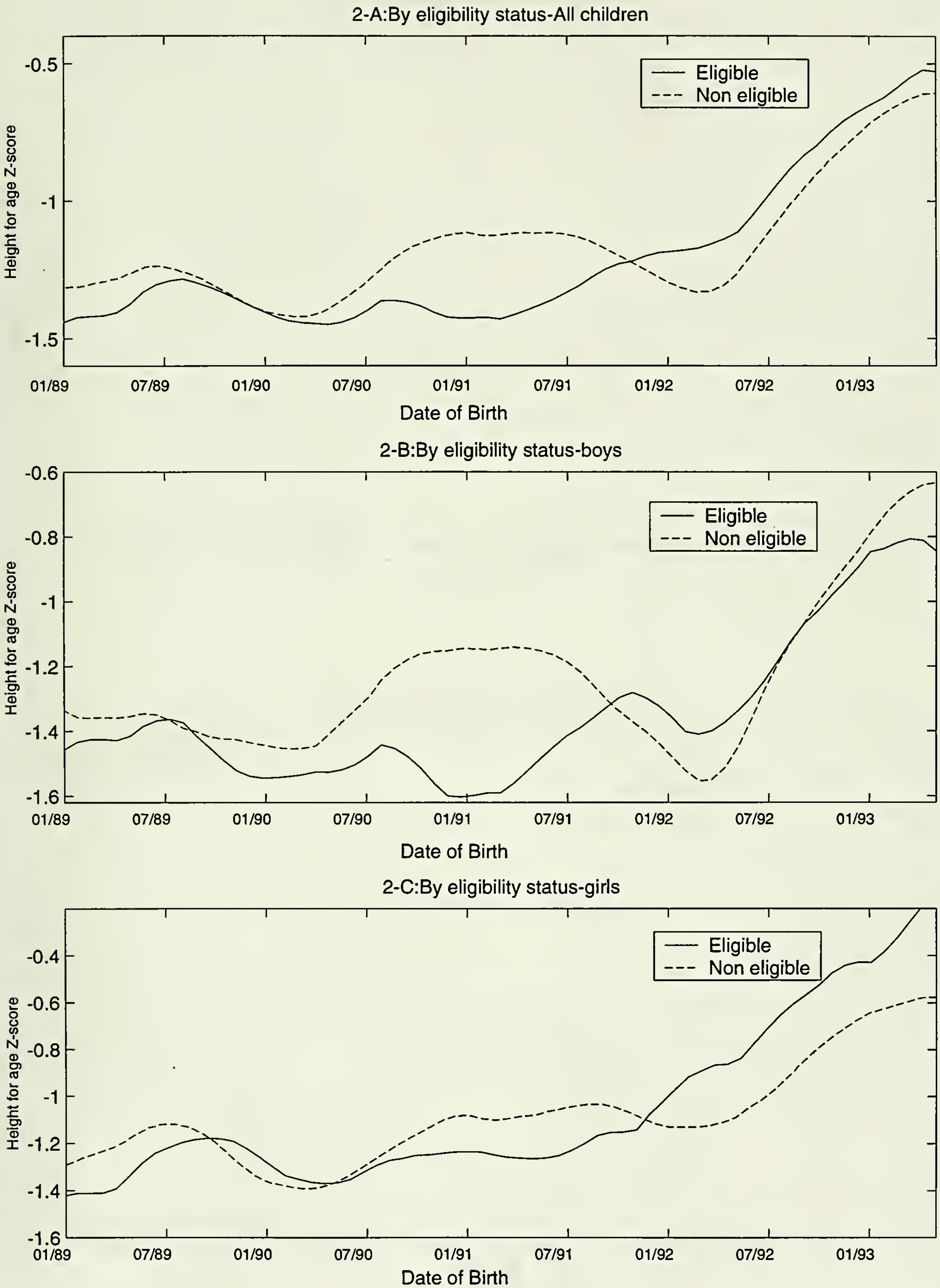
Note: Standard errors (robust to auto-correlation of residual within households) in parentheses.
Instruments in column 4 are indicator variables for woman and man eligible interacted with a dummy for child belonging to the youngest cohort. All equations include the uninteracted variables, year of birth dummies, family background variables, and family background variables interacted with year of birh dummies.

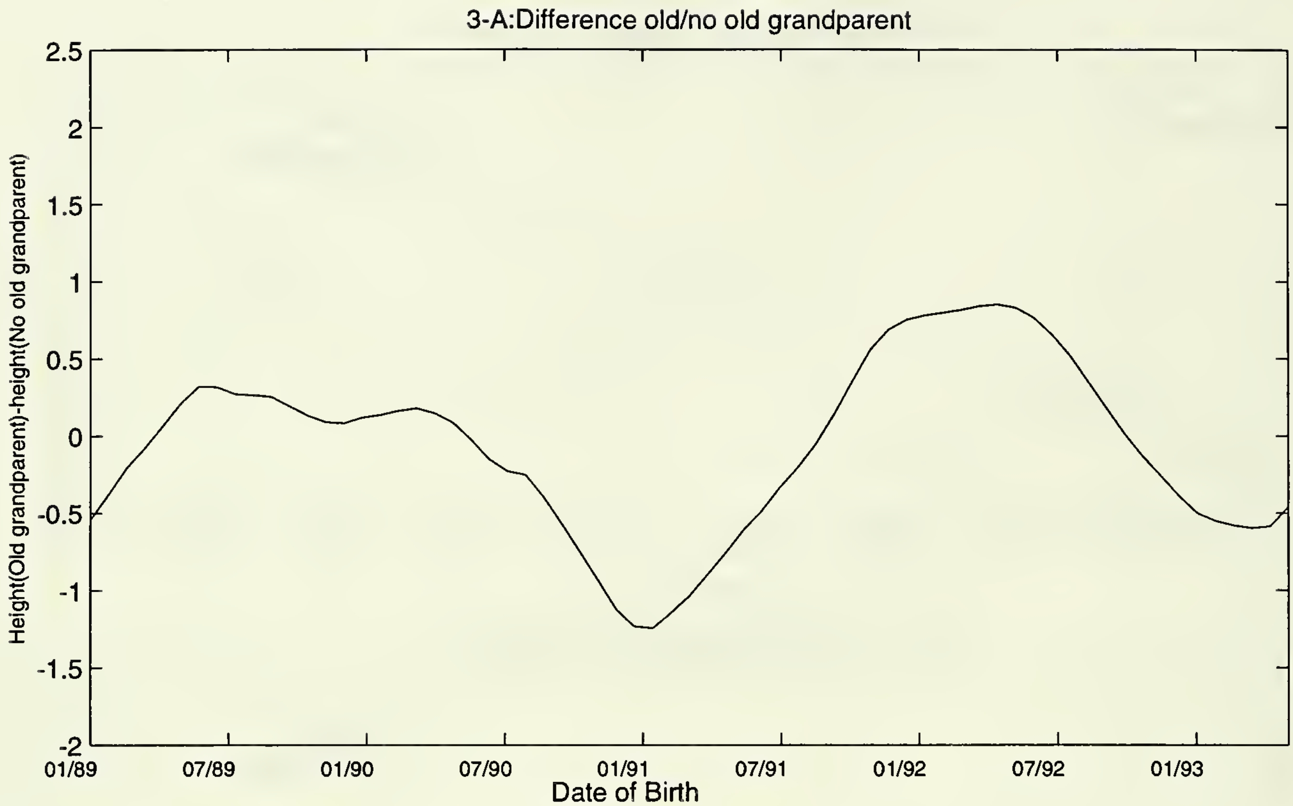
Table 12: Disposition of income

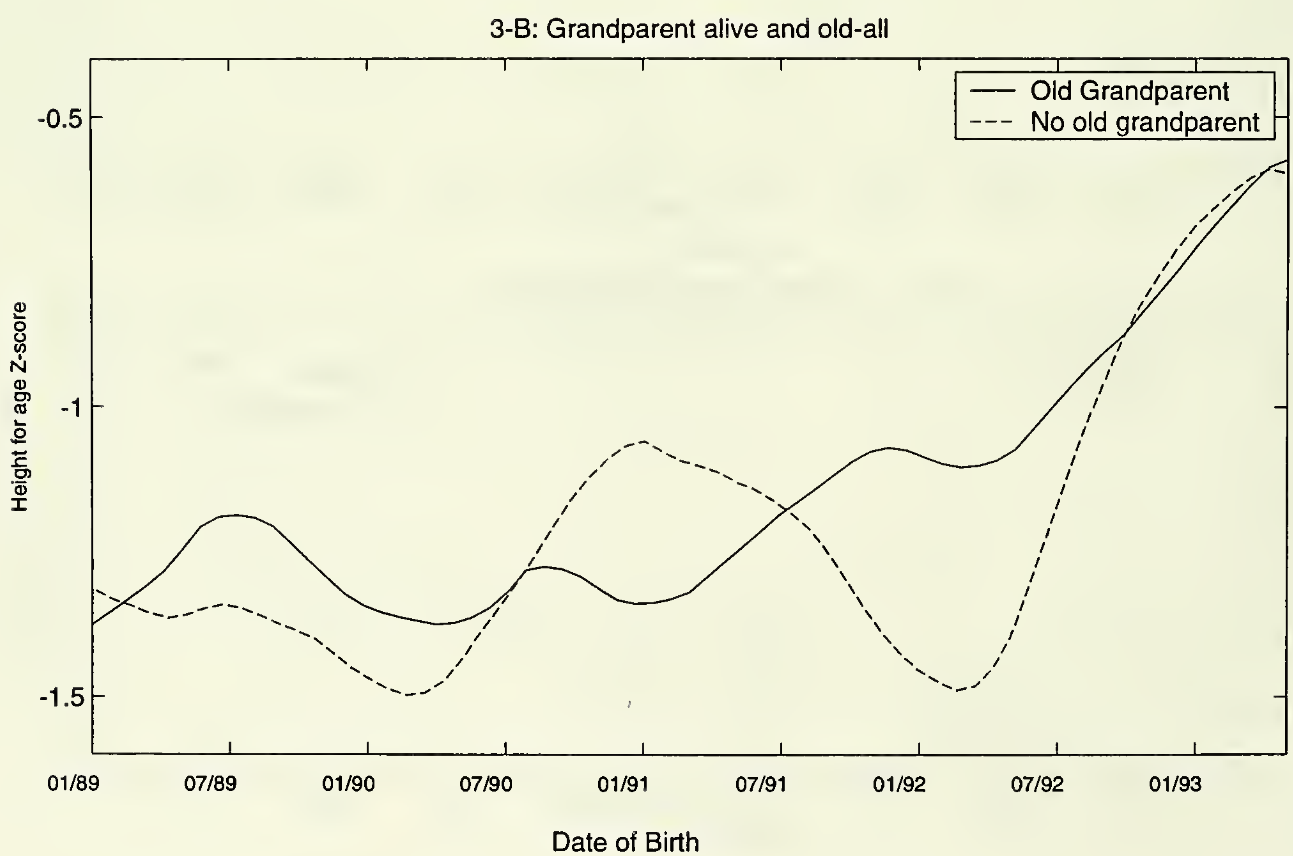
	Savings (formal+residual)	
	OLS	2SLS
	(1)	(2)
Woman's pension	0.99	0.82
income	(0.093)	(0.16)
Man's pension	0.78	0.53
income	(0.13)	(0.22)
Non-pension income	0.53	0.50
	(0.017)	(0.041)
F. test woman=man	1.49	1.1
(p. value)	(0.22)	(0.29)

Notes: Standard errors in parentheses.
Instruments are : dummies for: head is employed, head holds a regular job, a casual wage job, a job an agricultue, sector of the job, employer's type (central or local government, private firm, other), pay type (weekly, fortnightly, monthly), woman eligible and man eligible.

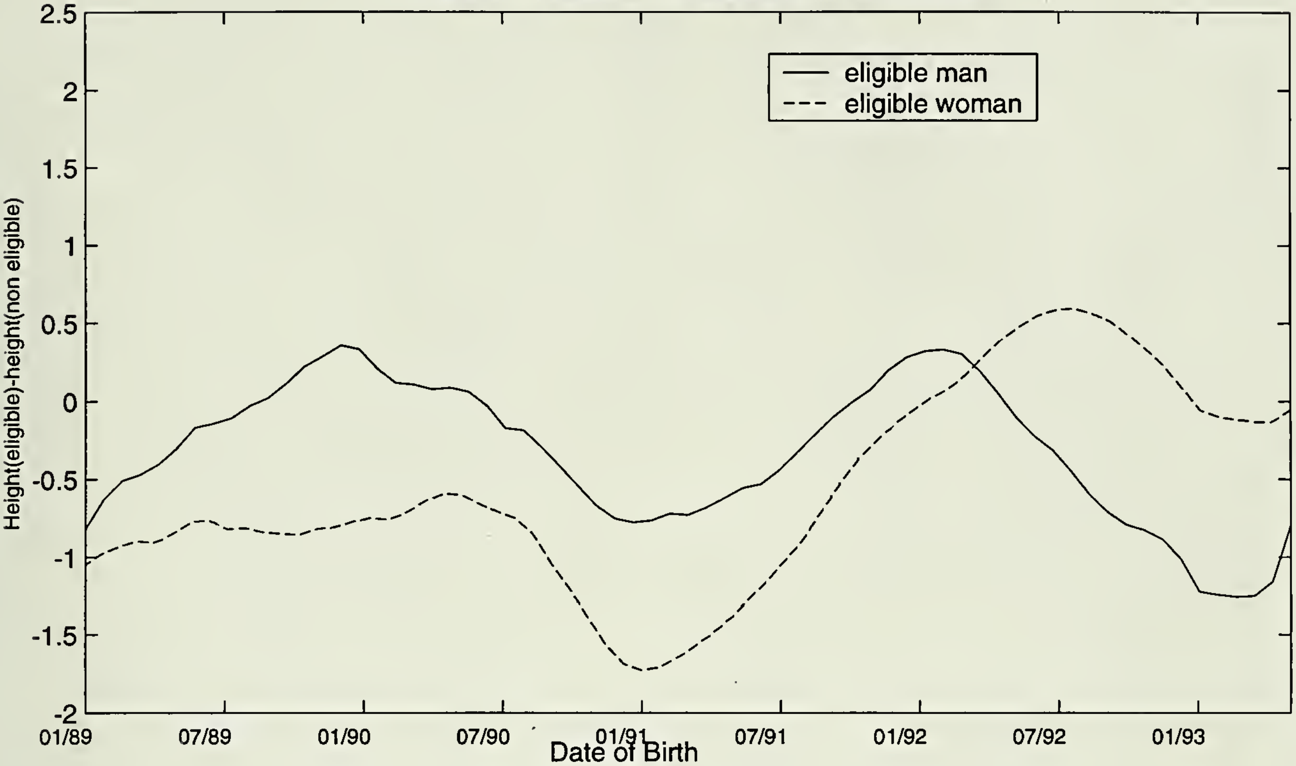

1-A:Height of eligible and non eligible children
Eligible
Non eligible
Height(centimeters)
105
100
95
90
85
80
75
70
65
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth
1-B:Difference eligible/non eligible (all children)
Height(eligible)-height(non eligible) (centimeters)
2.5
2
1.5
1
0.5
0
-0.5
-1
-1.5
-2
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth
1-C: Difference eligible/non eligible (boys and girls)
Boys
Girls
Height(eligible)-height(non eligible)
2.5
2
1.5
1
0.5
0
-0.5
-1
-1.5
-2
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth


2-A:By eligibility status-All children
Eligible
Non eligible
Height for age Z-score
-0.5
-1
-1.5
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth
2-B:By eligibility status-boys
Eligible
Non eligible
Height for age Z-score
-0.6
-0.8
-1
-1.2
-1.4
-1.6
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth
2-C:By eligibility status-girls
Eligible
Non eligible
Height for age Z-score
-0.4
-0.6
-0.8
-1
-1.2
-1.4
-1.6
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth


3-A:Difference old/no old grandparent
Height(Old grandparent)-height(No old grandparent)
2.5
2
1.5
1
0.5
0
-0.5
-1
-1.5
-2
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth


3-B: Grandparent alive and old-all
Old Grandparent
No old grandparent
Height for age Z-score
-0.5
-1
-1.5
01/89
07/89
01/90
07/90
01/91
07/91
01/92
07/92
01/93
Date of Birth

4-A:Difference eligible/non eligible (men and women)



4-B: Eligible women and eligible men

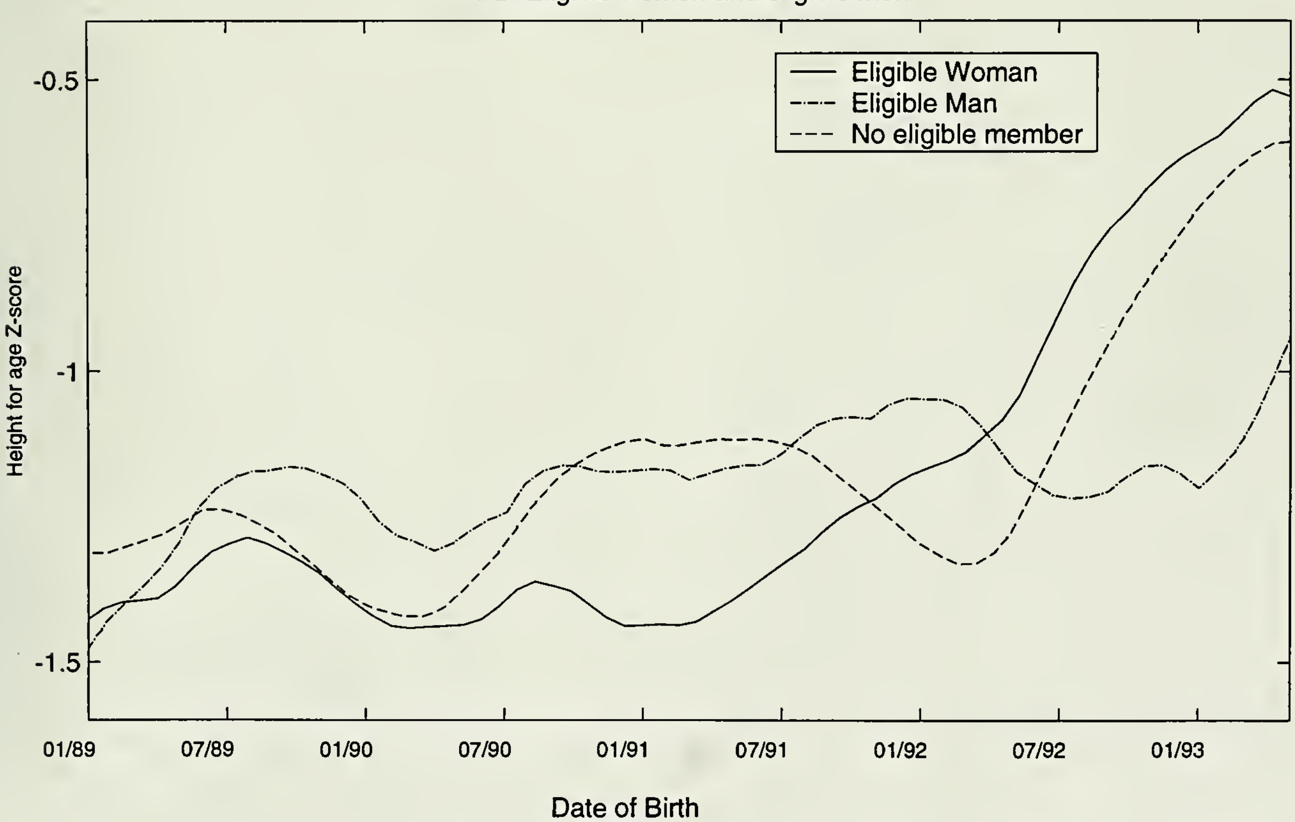


Table A1: Effect of eligibility for pension in various subsamples

	Household non-pension income per capita		Mother's education		Number of children aged less than 16		Number of eligible members		Age of eligible member	
	<median	>median	<=6 years	>6 years	>=4	<4	1	2 or more	>70	<=70
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
PANEL A: Boys										
A.1 Without controlling for presence of the grandparent in the household										
Eligible*born after 01/92	0.094	0.014	-0.32	0.44	-0.22	0.36	0.21	-0.66	-0.31	0.18
	(0.38)	(0.37)	(0.40)	(0.37)	(0.33)	(0.44)	(0.29)	(0.46)	(0.38)	(0.31)
Number of observations	635	512	717	555	721	551	1272		1272	
PANEL B: Girls										
B.1 Without controlling for presence of the grandparent in the household										
Eligible*born after 01/92	0.85	0.39	0.68	0.40	0.23	0.64	0.42	0.54	-0.097	0.52
	(0.43)	(0.39)	(0.43)	(0.35)	(0.35)	(0.47)	(0.29)	(0.59)	(0.43)	(0.33)
Number of observations	608	602	694	540	690	544	1234		1234	

Note: Standard errors (robust to auto-correlation of residual within households) in parentheses.
Year of birth dummies, family background variables and family background variables interacted with year of birth dummies are used as control variables in all regressions. The coefficients in columns 7 and 8 are obtained by rungin a single regression with one eligible and 2 eligible members or more entered separately.
The coefficients in columns 9 and 10 are obtained by running a single regression with young eligible and old eligible entered separately.